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# International trade and firm activity in an insecure world

Matteo Neri Lainé

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**DE L'UNIVERSITÉ PSL**

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**International Trade and Firm Activity in an  
Insecure World**

Soutenue par

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# International Trade and Firm Activity in an Insecure World

Matteo Neri--Lainé

*"Les principaux fondements que doivent avoir tous les Etats, les vieux comme les nouveaux ou les mixtes, sont les bonnes lois et les bonnes armes ; et puisqu'il ne peut y avoir de bonnes lois là où il n'y a pas de bonnes armes, et que là où il y a de bonnes armes il faut bien qu'il y ait de bonnes lois, je laisserai de côté les propos sur les lois et je parlerai des armes."*

Niccolò di Bernardo dei Machiavelli – *Le Prince*

---

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*Dans le chaos, un prince s'épanouit. L'iris rouge, il attend.*

*Une vérité sans morale. L'effort d'une pensée.*

*Sortez – de moi – chaque mot.*

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# Résumé en Français

Les activités économiques et le commerce sont étroitement liés à la gestion de la violence. Le concept central derrière cette relation est l'insécurité, définie comme un coût de production ou d'échange du fait d'une exposition à la violence. Pour y faire face, les Etats ont développé des politiques spécifiques. Mais, malgré les bénéfices économiques potentiels, la coopération internationale reste limitée. Cette situation questionne l'efficacité de ces politiques sécuritaires communes et les intérêts économiques des Etats à s'y engager. Cette thèse vise à apporter des éléments de réponse empiriques et théorique. Ainsi, elle se concentre plus particulièrement sur les coûts économiques de l'insécurité, la réponse des Etats en termes de politique et de coopération, et les différences de sensibilité aux coûts à l'échange.

Le premier chapitre étudie l'effet sur le commerce de traités conçus pour prévenir l'insécurité internationale : les alliances militaires. Mobilisant une approche basée sur la gravité structurelle, nous montrons que les alliances accroissent en moyenne le commerce de 60%. Les effets des alliances militaires sont dynamiques et hétérogènes. Ils dépendent étroitement du type d'alliance et la taille économique des partenaires. Une stratégie à variable instrumentale ainsi qu'une étude d'événement confirment l'interprétation causale de nos résultats. M'intéressant au mécanisme sous-jacent à l'effet des alliances militaires, je démontre que ces dernières accroissent les échanges commerciaux via la réduction de l'insécurité internationale. De plus, une analyse en équilibre général révèle que la croissance des flux commerciaux générée par les alliances militaires apporte des gains substantiels de bien-être pour les pays signataires, tandis que les pays non-alignés subissent des pertes.

Le second chapitre étudie l'effet de l'insécurité sur l'existence des entreprises formelles. Nous développons un cadre théorique flexible dans lequel l'insécurité affecte les décisions d'entrer sur le marché, de sortie et de formalisation des entreprises. Dans l'analyse empirique, je combine un ensemble de données original sur les entreprises afghanes avec des données géoréférencées sur les événements militaires survenus pendant le conflit afghan sur la période 2003-2016. Ainsi, je montre, que, dans un tel contexte de construction de l'État, l'exposition aux événements militaires a un effet positif moyen sur l'existence des entreprises formelles. Néanmoins, cet effet est très hétérogène selon les acteurs, l'emplacement, le moment et les caractéristiques des entreprises. Le conflit afghan a la particularité



d'impliquer profondément des pays étrangers. En mobilisant cette source particulière de variation exogène, j'identifie l'effet causal de l'insécurité sur l'existence des entreprises formelles. Je montre qu'une augmentation de 1% de l'exposition aux événements militaires instrumentés augmente la probabilité d'activité formelle de 4,16 %.

Dans le troisième chapitre, je développe un cadre théorique de commerce international et effectue une analyse de bien-être en présence d'élasticité commerciales spécifiques aux paires exportateur-importateur. Adaptant les modèles traditionnels du commerce, je démontre qu'en présence d'élasticité bilatérales, la part des dépenses domestiques et l'élasticité des exportations par rapport à l'indice des prix à l'importation sont des statistiques suffisantes pour évaluer les gains de bien-être issus du commerce. Je propose une méthodologie d'évaluation des politiques publiques en équilibre général. En appliquant cette méthode aux accords de libre-échange, nous mettons en évidence les distorsions importantes engendrées par les élasticité commerciales bilatérales. Si ces distorsions sont négligées, elles peuvent conduire à des estimations sensiblement erronées de l'impact de la libéralisation des échanges.

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# Introduction

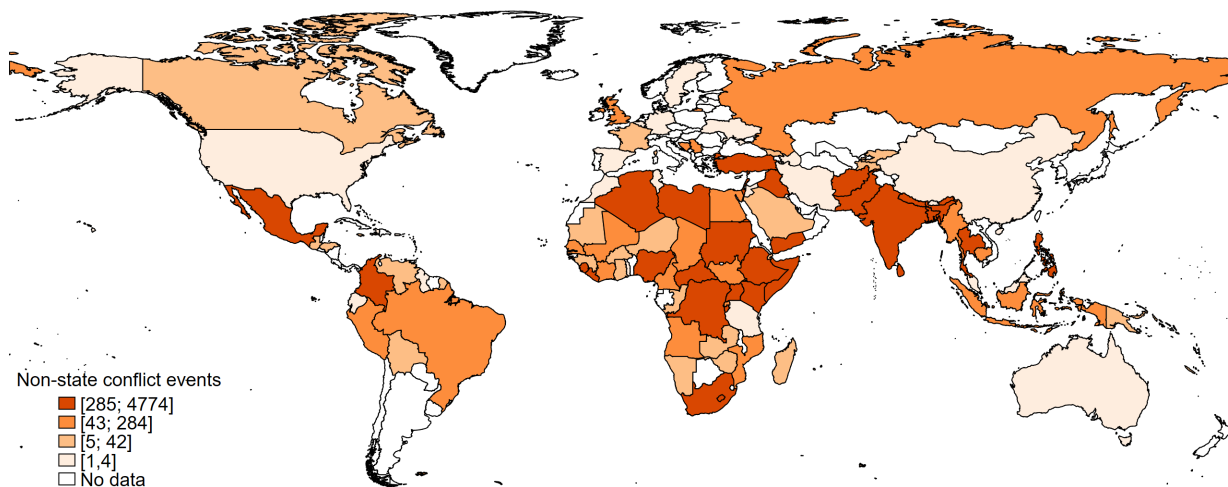
Economic activities and trade are inherently linked to the management of violence. In the modern global system, violence (i.e. expropriation, destruction, piracy, terrorism, conflicts, wars etc.) is still a prevalent and enduring challenge. No country is spared. Developing economies, characterized by significant weaknesses of central power and institutions, are constantly exposed to attacks of local armed factions vying for control over territories, resources and populations (cf. figure 1). In developed economies, direct state involvement in managing violence has succeeded in geographically expelling major threats from the national territory, but at the cost of a multiplication of military events and an internationalisation of security issues (cf. figure 1& 2).

This global exposure to violence imposes substantial economic costs. It leads to reductions in trade ranging from  $-30\%$  to  $-80\%$  (Anderson and Marcouiller, 2002; Blomberg and Hess, 2006; Glick and Taylor, 2010; Martin et al., 2008; Sandkamp et al., 2022; Couttenier et al., 2023). It severely impacts the activities of firms (Camacho and Rodriguez, 2013; Couttenier et al., 2022; Klapper et al., 2013; Del Prete et al., 2023; Verwimp et al., 2019) and more broadly, impedes economic growth and development (Gates et al., 2012; Koubi, 2005; Murdoch and Sandler, 2002, 2004; Sandler and Enders, 2008; Thies and Baum, 2020). The central concept underlying this negative relationship between violence and economic activities is insecurity, defined as a production or exchange cost stemming from exposure to violence (Anderson and Marcouiller, 2002; Martin et al., 2008; Couttenier et al., 2023; Del Prete et al., 2023).

States have understood the importance of reducing international insecurity and developed specific policies which may have important economic returns (Findlay and O'Rourke, 2009; Gibler, 2008; Mansfield and Bronson, 1997; Martin et al., 2008; Thoenig, 2023). Nonetheless, international cooperation is still limited (Sundberg and Melander, 2013), questioning its efficiency and the economic interest of countries in expanding it. This thesis attempts to provide empirical and theoretical answers. The first chapter focuses on the effect of military alliances on trade – treaties specifically designed to reduce international insecurity. In the second chapter, we dig into the complexity of military events and investigate their micro-economic impact on formal firms. Finally, the third chapter takes a more

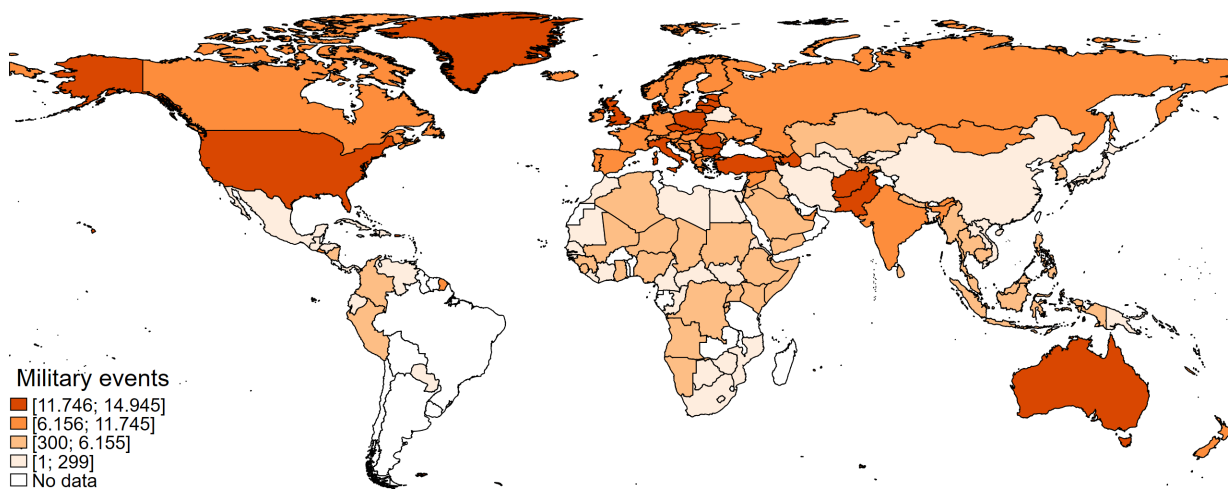
general perspective and analyses the consequences of country-pair-specific sensitivity to exchange costs on real-revenues. The following of this introduction explains the main contributions and takeaways of each chapter.

Figure 1: Countries' exposure to non-state conflict events



Note: Non-state conflict events are incidents where armed force was used by a non-state organised actor against another organised actor or civilians, resulting in at least one direct death at a specific location and a specific date. We display the total number of events countries' have been exposed from 1989 to 2018. Data are from the geocoded Uppsala Conflict Data Program (UCDP) project (Sundberg and Melander, 2013) Maps' categories are defined by quartiles.

Figure 2: Countries' exposure to military events



Note: Military events are incidents where armed force was used by an organised actor against another organised actor, whose at least one of them is a state, resulting in at least one direct death at a specific location and a specific date. We display the total number of events countries' have been exposed from 1989 to 2018. Data are from the geocoded Uppsala Conflict Data Program (UCDP) project (Sundberg and Melander, 2013) Maps' categories are defined by quartiles.

**Chapter 1** – This chapter explores a policy for reducing trade insecurity: military alliances. These international agreements are designed specifically to decrease insecurity among their members, based on two main principles: (1) enforcement of military cooperation policies, and (2) international security

as a way to promote trade.

Taking a structural gravity approach, I estimate the effects of alliances on bilateral exports. On average, enforcing such agreements increases trade by 60%, equivalent to a tariff reduction of 12.8%. However, this effect varies widely depending on the type of treaty and the economic size of the partners. Thus, only deep alliances have a significant effect, while small countries benefit more than large economies from their enforcement. Furthermore, an instrumental variable strategy and an event study confirm the causal interpretation of the results.

Subsequently, the chapter analyses the mechanism by which military alliances influence bilateral trade. Modern conflicts primarily involve combating non-state actors (terrorism, civil wars, piracy, armed factions). Retrieving data on conflict events, I construct an extensive measure of international insecurity. Employing a two-stage strategy, I demonstrate that alliances substantially reduce international insecurity, leading to a significant increase in bilateral exports.

Finally, the chapter studies the welfare gains from alliances. In a general equilibrium setting, I use the Pseudo-Poisson-Maximum-Likelihood estimator properties to construct a counterfactual for 2012 in which all alliances are cancelled. In this way, I assess the change in real revenues induced by the network of alliances. These treaties bring substantial welfare benefits to their members but also a welfare cost to neutral countries through a trade diversion mechanism. Additionally, through scenario analysis, I demonstrate the significant potential welfare implications of reshaping the military alliance network in response to the war in Ukraine.

**Chapter 2** – The second chapter studies the effect of insecurity on formal firms' existence. Formal economic activities play a crucial role in development, as they house the most productive firms and are essential for state finances. Their sensitivity to insecurity is therefore a primary concern. To investigate this question, conflict events are commonly used as a direct proxy for insecurity. Yet, conflicts are very heterogeneous (actors, objectives, duration, etc...), and, in certain contexts, conflict events may improve security through the development of the state capacity. Working on the 2003-2016 Afghan conflict, the study has the pertinent state-building context to dig into the complexity of conflict events and discuss the consequences of insecurity.

Firstly, I present a theoretical model where heterogeneous firms endogenously enter the market, exit, and choose their (in)formality status. In this framework, I exhibit how insecurity affects firms' choices and derive the insecurity elasticity in the formal activity probability. The model predicts that a reduction of 1% in insecurity increases the formal firms' existence between 7.86% and 11.86%.

I conduct an empirical analysis. For this purpose, I use panel data from the former Afghan Ministry of the Economy's register of formal companies and geocoded observations of military events. Based on the theoretical expression of the formal activity probability, I demonstrate that, on average, exposure



to military events has a slight positive effect on the existence of formal firms. In the context of Afghan state-building from 2003 to 2016, military events signal an expansion of state capacity, which, in turn, reduces insecurity for formal firms and increases the probability of formal activity. Exploring the heterogeneity of military events, I show that their impact heavily depends on factors such as actors, timing, and location.

The post-2003 Afghan conflict is characterized by significant involvement from external countries. Leveraging this source of exogenous variation, I adopt an instrumental variable strategy. I estimate the impact of changes in military events determined by the geographic allocation (share) of exogenous variations in NATO's military events (shift). This approach allows us to address endogeneity issues, isolate military events signalling a state capacity expansion, and identify a clear change in insecurity. The findings indicate that a 1% increase in exposure to instrumented military events raises the probability of formal activity by 4.16%. Additionally, we uncover micro-level heterogeneities in insecurity reduction, favouring small firms and those with a NATO member nationality, which significantly distorts the supply structure.

**Chapter 3** – In classical trade models used to perform welfare analysis, the trade elasticity is the key parameter ([Arkolakis et al., 2012](#)). Traditionally assumed to be constant, recent empirical studies have provided evidence of exporter-importer-specific trade elasticities consistent with modern structural gravity estimations. This empirical discovery of bilateral elasticities prompts the need to comprehend how gains from trade or any reduction in exchange costs should be assessed.

In the third chapter, I develop a theoretical framework for international trade and conduct a welfare analysis considering exporter-importer-specific trade elasticities. I introduce bilateral trade elasticities into two categories of models fundamental for modern trade welfare analysis: Armington models ([Anderson, 1979](#); [Anderson and Van Wincoop, 2003](#); [Armington, 1969](#)) and heterogeneous firms models ([Chaney, 2008](#); [Melitz, 2003](#)). I discuss the changes in micro-foundation assumptions, their theoretical implications, and the resulting structural gravity equations.

The first outcome of the chapter is that, in the presence of bilateral trade elasticities, the share of domestic expenditure and the elasticity of exports to the importer's price index (referred to as price-index elasticity) suffice to infer welfare gains from trade. Employing a Method of Moments Quantile Regression, I estimate bilateral trade elasticities and derive price-index elasticities. By computing welfare gains from trade, I demonstrate that constant trade elasticities tend to overestimate gains for small economies and underestimate them for large economies.

Then, in the second part of the chapter, I propose a methodology to perform trade policy evaluations with bilateral trade elasticities. Using the Poisson Pseudo Maximum Likelihood (PPML) estimator properties, I adapt the methodology of [Anderson and Yotov \(2016\)](#) and its extension in [Anderson et al.](#)

(2018). Distortions in trade, price index, and wage elasticities lead to welfare outcomes significantly different from those obtained with constant trade elasticities. Using my method to evaluate the impact of Regional Trade Agreements (RTAs), I demonstrate that the constant trade elasticity approach tends to overestimate welfare gains.

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# Chapter 1

## Sovereign Gravity

### The Military Alliance Effect on Trade

International insecurity can severely disrupt trade. This chapter studies treaties aimed at preventing such insecurity: military alliances. Taking a structural gravity approach, we show that alliances increase trade by 60% on average. The effects of military alliances are dynamic and heterogeneous. They depend to a large extent on the type of alliance and the economic size of partners. An instrumental variable strategy and an event study confirm the causal interpretation of the results. Investigating the mechanism behind the impacts of military alliances, we demonstrate that alliances increase trade by reducing international insecurity. General equilibrium analysis moreover shows that the growth in trade generated by military alliances brings substantial welfare gains for signatories and losses for non-aligned countries.

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Since 2000, the number of major conflicts has quadrupled worldwide to involve more than 130 countries.<sup>1</sup> This massive increase in international insecurity can severely disrupt trade. The vast majority of trade costs are not associated with direct policy instruments, but with hidden transaction costs (Anderson and Van Wincoop, 2004). A significant proportion of these hidden transaction costs have to do with the insecurity of trade (Anderson and Marcouiller, 2002). Blomberg and Hess (2006) show that common exposure to violence reduces bilateral trade by 7%, a figure that increases to 35% in the case of civil war. Sandkamp et al. (2022) determine that each additional maritime piracy incident reduces bilateral exports by 0.1%. Korovkin and Makarin (2023); Rohner et al. (2013); Yu et al. (2015) demonstrate the importance of conflict signals in shaping international trade costs. Martin et al. (2008) show that military interstate disputes reduce bilateral trade by 38% on average. Glick and Taylor (2010) demonstrate the overall negative impact of war on trade. They show that major interstate conflicts reduce trade by 80% between enemies and by 13% between belligerents and neutral countries, with a significant lasting effect in peacetime. Introducing conflict risk into a conventional general equilibrium model of trade, Thoenig (2023) exhibits the major interplays between interstate wars, diplomatic negotiations and trade.

This chapter examines one way of reducing trade insecurity: *the military alliances*. These international agreements are specifically designed to reduce insecurity among their members. They are based on two pillars: (1) enforcement of military cooperation policies, and (2) international security as a way to promote trade.<sup>2</sup> Many alliances exist such as the North Atlantic Treaty Organization (NATO), the Treaty on Collective Security and the Arab-Maghreb Union. However, the *Pax Mongolica* is an iconic example of such an alliance. In the 13th and 14th centuries, this set of treaties<sup>3</sup> ensured the security and development of trade in Eurasia. The end of the agreement saw a huge increase in conflicts and a sharp drop in trade – enough to prompt Europeans to take an unprecedented step in search of new trade roads (Findlay and O’Rourke, 2009). There are basically two categories of alliance: weak alliances – reducing the probability of open conflict between signatories – and defence pacts – enforcing collective and centralised management of members’ security (Gibler, 2008). Their depth of military cooperation is significantly different, as are their expected effects on insecurity and trade.

Few previous papers have analysed the impacts of alliances on trade. Those that do are restricted

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<sup>1</sup>Authors calculation based on geocoded UCDP project data on conflict events (Sundberg and Melander, 2013).

<sup>2</sup>“An alliance is a formal contingent commitment by two or more states to some future action. The action involved could entail almost anything—detailed military planning, consultation during a crisis, or a promise by one state to abstain from an upcoming war. [...] empirical studies have developed a consensus that the operationalisation of the alliance variable depends on two factors. First, alliance members have to be independent nation-members of the international system (for example, so-called alliances between international terrorist organisations do not qualify), and second, a treaty text has to exist that identifies a military commitment that is defensive, a neutrality arrangement, or an “understanding” such as an entente,” (Gibler, 2008). Details on alliances’ content are provided in section 1.A.

<sup>3</sup>The Pax Mongolica was a set of treaties between the former Mongol empire states – The Golden Horde (Western Steppe), the Yuan Empire (China), the Ilkhanat (Persia) and the Chagatai Khanate (Eastern Steppe) – the Italian republics and the Russian duchies (Findlay and O’Rourke, 2009).

to the Cold War period, find heterogeneous results and lack theoretical grounding. Drawing on the [Tinbergen \(1962\)](#) gravity equation, [Mansfield and Bronson \(1997\)](#) find a positive correlation between exports and alliances using a panel regression covering the 1960-1990 period. They conclude that an alliance increases exports by 20%. Taking a similar approach, [Long \(2003\)](#) estimates for the 1885-1990 period that defence pacts are associated with 37% higher exports, while weak alliances have no statistically significant effect.

Our study of the impact of military alliances on trade contributes to the literature in a number of ways. First, we identify the causal effect of alliances. On average, military alliances increase bilateral exports by 60%. Second, we investigate the mechanism behind this impact to show that alliances boost trade by significantly reducing international insecurity. Third, to the best of our knowledge, we are the first to investigate the welfare effect of the growth in trade induced by alliances. We show that the enforcement of an alliance brings substantial welfare gains for signatories, but losses for non-aligned countries.

This study is structured as follows. We present the focus on military alliances under structural gravity theory. We isolate the costs of insecurity in a model covering heterogeneous firms based on [Chaney \(2008\)](#) and [Helpman et al. \(2008\)](#). By reducing trade costs sensitive to insecurity, military alliances directly increase trade between partners. From this frame, we derive the gravity equation on which our empirical work is based. The analysis combines the Correlate of War database, built on the massive and meticulous work by [Gibler \(2008\)](#) to document active military alliances, with the CEPII<sup>4</sup> CHELEM dataset on international trade. Thus, we perform a structural gravity approach using a panel of 6,972 country pairs covering the period from 1967 to 2012.

We estimate the effects of military alliances on bilateral exports. Taking exporter-year, importer-year and exporter-importer fixed effects, our specifications focus on the within-country-pair variation of military alliances. Our set of fixed effects ensures that we properly control for multilateral resistance terms, market access and structural interstate relationships ([Behrens et al., 2012](#); [Feenstra, 2015](#); [Redding and Venables, 2004](#)). In addition, we control for Regional Trade Agreements (RTAs). The spread of RTAs and alliances over our period is highly distinct.<sup>5</sup> Yet, since we are investigating the specific effect of alliances, it is important to control for the standard agreements designed to affect trade. On average, enforcing a military alliance increases bilateral exports by 60%, which is equivalent to a tariff reduction of 12.8%.<sup>6</sup> This result is robust to a wide range of consistency checks, including additional controls (tariffs, depth of RTAs, Cold War, etc.), but also to other estimation techniques preventing potential bias (intranational trade, negative weights, asymptotic bias, etc.). Nonetheless,

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<sup>4</sup>Centre d'Etudes Prospectives et d'Informations Internationales.

<sup>5</sup>See graph [1.G.1](#) in appendices.

<sup>6</sup>Equivalence is made with the estimated trade elasticity in our sample  $\theta = 3.7$ .



the effects of alliances are highly heterogeneous. They are sensitive to the nature of the treaty and the economic size of partners. Thus, only defence pacts have a significant effect, while small countries benefit more than large economies from the enforcement of such agreements.

We carefully investigate the endogeneity of military alliances. Using an instrumental variable strategy supported by a plausible exogeneity test based on [Conley et al. \(2012\)](#), we confirm the causal interpretation of our results. We give a particular attention to the dynamic effect of alliances. Alliances signatures may be driven by past changes in international relations and require time to be fully enforced. Developing a Differenced Average Treatment on the Treated (DATTT) approach based on [Couch and Placzek \(2010\)](#), we address the pre-trend issue and show that the signature of an alliance has an immediate positive effect on bilateral trade but needs four years to reach their maximum effect and stabilise.

Then, we turn to analysing the mechanism by which military alliances affect bilateral trade. We directly test whether alliances increase trade by reducing international insecurity. The first source of insecurity that alliances address is the risk of conflict between members. However, we show that this channel does not apply. Indeed, modern conflicts are in the extremely large majority related to fights against non-state actors (terrorism, civil wars, piracy, armed factions). Retrieving data on conflict events from the geocoded UCDP project, we construct an extensive measure of country-pairs' insecurity including conflicts with non-state actors. Using a two-stage strategy and considering the heterogeneity of alliance treaties, we show that: (i) defence pacts sharply reduce bilateral insecurity, and (ii) by reducing insecurity, they significantly increase bilateral exports. This explains the effects of defence pacts on trade as a whole. Therefore, our results strongly support both the validity and prevalence of the insecurity mechanism.

In the last part of the chapter, we investigate the welfare effect of alliances. We develop a general equilibrium analysis. In keeping with [Arkolakis et al. \(2012\)](#), we derive the welfare system from our theoretical model, pointing up the role of insecurity costs. Then, using the properties of the PPML estimator, we solve this system and perform a counterfactual analysis for 2012 in which all alliances are ended. This enables us to draw conclusions about the impacts of military alliances on real revenue – our measure of welfare. Military alliances bring their members substantial welfare gains. Interestingly, our results show that neutral countries experience a marked welfare cost at the same time. Moreover, performing a scenario analysis, we show the considerable potential welfare ramifications of reshaping the military alliance network in response to the war in Ukraine.

The chapter is organised as follows. Section 1 describes the theoretical framework. Section 2 presents the data used in the analysis and some descriptive evidence. Section 3 investigates the effects of military alliances on bilateral exports, the sensitivity of our baseline results and heterogeneous

effects. Section 4 addresses potential endogeneity concerns. Section 5 investigates the dynamic effect of alliances. Section 6 studies the mechanism through which alliances affect trade. Section 7 develops the general equilibrium analysis and draws conclusions about the welfare impact of military alliances. Lastly, section 8 presents a short conclusion.

## 1.1 Theoretical framework: Insecurity and structural gravity

When two countries sign an alliance, they enforce bilateral security cooperation policies (Gibler, 2008) – see appendix 1.A for details and examples. In so doing, they improve security between partners. In other words, by reducing insecurity costs, alliances are supposed to increase trade. Below, we present the focus on military alliances under structural gravity theory and the resulting gravity equation. The full model and its extensions with fixed insecurity costs and bilateral trade-elasticities are detailed in the appendices (Section 1.B).

We define insecurity as a variable cost induced by exposure to violence (destruction, war, conflicts, terrorism, piracy). As discussed by Martin et al. (2008), the insecurity cost can be either country-specific or country-pair-specific. Alliances are international treaties which focus on reducing the probability of conflict between the members and developing security cooperation policies for common purposes. They do not aim to reduce the global insecurity of a country, but the insecurity which concerns signatories (see appendix section 1.A). Signing an alliance should, therefore, only affect the country-pair insecurity cost.<sup>7</sup>

From the perspective of a firm, the country-pair insecurity cost is interpretable as the probability of a product to be expropriated when sold in country  $j$  from country  $i$ . When a firm exports in the presence of insecurity, each product has a probability  $S_{ij}$  of being sold and  $1 - S_{ij}$  of being expropriated. Firms export a number of units of their variety, so the probability of expropriation is also interpretable as the share of exports that does not arrive at destination (i.e. the consumer). This is similar to a traditional iceberg trade cost  $\tau$ , but sensitive to insecurity with  $\tau_{ij}(s) = 1/S_{ij}$ .<sup>8</sup> In monopolistic competition with a Constant Elasticity of Substitution (CES) demand function, this leads to the price function:

$$p_{ij}(\alpha) = \frac{\sigma}{\sigma - 1} w_i T_{n,ij} \alpha \quad (1.1)$$

<sup>7</sup>Following this theoretical discussion, the empirical sections of the chapter focuses on the bilateral effect of alliances. Yet we further discuss the potential role of multilateral insecurity in the appendix section 1.C.

<sup>8</sup>We can also interpret the reduction of the expropriation probability as a lower cost of insurance. In the case of an insurance market, a firm can pay insurance which, in exchange for a contribution equal to the share  $\tau_{In,ij}$  of the value of each insured exported product, will provide the amount  $p$  (the price) for each expropriated product. Thus, in exchange for a variable cost  $\tau_{In,ij}$ , the firm obtains the guarantee that the exported products will be sold at the price  $p$ . Given that we are in a Melitz (2003) monopolistic competition case, firms face a returns-to-scale technology due to the presence of fixed costs and are price setters. Therefore the firm chooses the lower price between  $p_{ij}(\alpha) = \frac{\sigma}{\sigma - 1} w_i \tau_{s,ij} \alpha$  and  $p_{ij}(\alpha) = \frac{\sigma}{\sigma - 1} w_i \tau_{In,ij} \alpha$ . Firms only take out insurance policies that will ensure  $\tau_{In,ij} \leq \tau_{s,ij}$ . So, when the probability of expropriation decreases, there is a reduction in  $\tau_{In,ij}$ . Even in the presence of an insurance market, military alliances reduce bilateral insecurity costs.

where  $\sigma$  is the elasticity of substitution,  $w_i$  the wage in country  $i$ ,  $\alpha$  the firm's marginal cost (i.e. the inverse of productivity  $\gamma$ ).  $T_{n,ij} = \Pi^n \tau_{n,ij}$  is a product of variable trade costs with  $n$  the  $n$  potential source of iceberg cost, including  $\tau_{s,ij}$  the variable insecurity cost, but also all variable trade costs sensitive to other parameters (geography, standard trade policies, institutions, etc.). Thus, in such a frame, any reduction in bilateral insecurity reduces exporting firms' prices.

From this theoretical frame, we can derive at the aggregated level the following structural gravity equation outlying the insecurity trade costs  $\tau_{s,ij}$ :

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} T_{n \neq s, ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.2)$$

$X_{ij}$  is the total exports from country  $i$  to country  $j$ ,  $N_i$ , the number of firms in the exporting country,  $\bar{\alpha}_i$ , the maximum marginal cost of country  $i$ 's technology, and  $w_i$  the wage in country  $i$ 's economy.  $X_j$  is the total revenue of country  $j$ , and  $\Phi_j$ , the importer's multilateral resistance term, while  $T_{n \neq s, ij}$  and  $F_{n \neq s, ij}$  are respectively the variable and fixed trade costs sets insensitive to insecurity.<sup>9</sup> Like other trade costs, insecurity costs ( $\tau_{s,ij}$ ) have a negative elasticity.<sup>10</sup> Hence, by reducing the insecurity costs, the enforcement of a military alliance between countries  $i$  and  $j$  increases bilateral exports  $X_{ij}$ .

## 1.2 Data

The structure of the dataset is a country-pair panel. Our unit of observation, therefore, is a given exporter-importer-year combination. We study how variations in the ally status of the dyad affect bilateral exports.

### 1.2.1 Data description

*Alliances data.*— We use information on military alliances for each  $ijt$  from the Correlate of War project (Gibler, 2008). We have information on whether a given country pair are allies and, if so, the nature of the treaty. We can divide military alliances into two categories: weak alliances, which focus mainly on military cooperation to guarantee peace between signatories, and defence pacts, which enforce military cooperation to protect members from common threats and achieve common strategic objectives<sup>11</sup> (Long, 2003; Gibler, 2008). From 1967 to 2012, the majority of military alliances were defence pacts. Yet, most of them were enforced throughout the entire period. Consequently, in our sample the number of defence pacts and weak alliances contributing to the within variation – i.e.

<sup>9</sup>See section 1.B.2 for the model extension with fixed insecurity costs.

<sup>10</sup>Note also that despite our assumption that all trade costs of the same nature (variable vs fixed) have the same elasticity, this does not mean that the model assumes that all policies have the same trade elasticity. Indeed, we do not assume that all trade costs are sensitive to the same policies or to the same extent.

<sup>11</sup>See section 1.A for more information about military alliance treaties and examples.

whose status changes over the period – is comparable (see table 1.1). For each country pair, we define  $ALL_{ijt}$ , a dummy variable which equals 1 if country  $i$  and  $j$  are allies at time  $t$  and 0 otherwise.

*Insecurity and military cooperation data.*– We collect data on conflict events from the geocoded UCDP project (Sundberg and Melander, 2013) to construct our measures of military cooperation and bilateral insecurity. The initial observation unit is an event. Information is available starting in 1989 with the year provided for each event. This project also has the advantage of identifying the belligerents (and co-belligerents) in each conflict event. In the large majority, observed events are related to conflicts between states and non-state organized actors (terrorism, civil wars, piracy, armed factions). We organise the information to create a dummy taking the value one if the country-pair cooperates militarily (i.e. is belligerent) in the event. Summing this dummy at the country-pair-year level, we obtain a continuous measure of bilateral military cooperation. By summing observations of conflict events, excluding the country-pair’s cooperative events, we observe country-year exposure to insecurity.

*Exports data.*– International trade data are retrieved from the CEPII CHELEM base (de Saint-Vaulry, 2008). We extract bilateral exports in current dollars between the 84 available countries from 1967 to 2012.<sup>12</sup> Data exclude re-imports and re-exports. Flows are adjusted for freight and insurance costs.<sup>13</sup> Zero trade flows are observed and our export matrix is squared. The CHELEM database enables us to exclude arms exports, which improves our identification strategy. Therefore, in our case, CHELEM provides the best trade-off between quality of observations and panel size. Nevertheless, we discuss below a robustness check with extended trade data. The CHELEM base does not provide information on intranational trade. We supplement our export data with within-country flows from the CEPII TradeProd database (De Sousa et al., 2012; Mayer et al., 2023).<sup>14</sup> Similarly to CHELEM, flows are adjusted for freight and insurance costs and harmonised according to the reliability of the countries’ declarations. However, the dataset starts in 1990 and 19 CHELEM sample countries are unobserved.<sup>15</sup> Within-country trade flows are very important for the general equilibrium analysis performed in the last part of the chapter. Yet, because of this data limitation, the baseline and the other partial equilibrium estimations use only international exports. In this way, we make use of the

<sup>12</sup>The full list of countries is presented in appendix table 1.G.1.

<sup>13</sup>In declarations, imports include freight and insurance costs while exports do not. Considering the reliability of countries’ declarations, CHELEM’s bilateral trade is harmonised in keeping with the RAS iterative method (see Stone (1963)). Prior to 1992-93, some countries, such as the former USSR and former Yugoslavia, are not recognised (or reported) as independent trade partners by the UN. The CHELEM base provides estimated values to fill these missing observations. The dataset therefore contains harmonised export values for all exporter-importer pairs (6,972) across the entire period.

<sup>14</sup>Intranational flows are filled by linear interpolation of non-missing data, whereas the remaining missing values are extrapolated using country total exports (Baier et al., 2019; Fontagné and Santoni, 2021).

<sup>15</sup>Excluded countries are: Belarus, Bosnia and Herzegovina, Brunei, Croatia, Czech Republic, Estonia, Kazakhstan, Kyrgyzstan, Latvia, Libya, Lithuania, Luxembourg, North Macedonia, Paraguay, Russia, Slovakia, Slovenia, Ukraine. A recent version of TradeProd includes data from 1966. Yet the covered period is very heterogeneous among countries especially before 1990 and after 2008. We, therefore, use the same TradeProd dataset as Fontagné and Santoni (2021) which is designed for general equilibrium analysis.

largest available panel. Nonetheless, a robustness check with intranational flows is discussed below.

*Other data.*— Information on Regional Trade Agreements (RTAs) and the standard gravity variable, such as distance, population, common language and religion, are retrieved from the CEPII’s Gravity database (Head et al., 2010). Data on RTAs include preferential trade agreements, free trade agreements, customs unions and other less common forms of agreements. We round them out with RTA legally enforceable provisions from the Content of Deep Trade Agreements database (Hofmann et al., 2017). Information on Gross Domestic Product (GDP) is also taken from the CHELEM dataset. Information on tariffs is retrieved from the World Trade Integrated Solution, which combines data from UNCTAD TRAINS<sup>16</sup> and the World Trade Organization. Finally, we extract our data on corruption and the rule of law from the Variety of Democracy project (Coppedge et al., 2022; Pemstein et al., 2022).

### 1.2.2 Descriptive statistics

We have a panel of 6,972 country-pairs from 1967 to 2012. In our dataset, 46% of worldwide exports (in value) come under the umbrella of alliances: 72 countries are signatories to these treaties and 738 pairs are affected by one, including 362 making a switch during our period (cf. table 1.1).

In figure 1.1, the map of the world displays the number of alliances per country during our time-frame. As can be observed, alliances are heterogeneously distributed across countries. No clear correlation between level of economic development and being signatories to such treaties is observed. Intermediary or low-income countries are not excluded from the worldwide alliance system – South American countries have signed more alliances than any European countries, while Africa and Asia present a wide range of involvement. In figure 1.2, the same exercise is replicated with the number of switches in alliance per country. Countries contributing to the switches are well dispersed around the globe, providing a good range of treated economies and international relationships.

Table 1.1: Alliances, treated and untreated countries

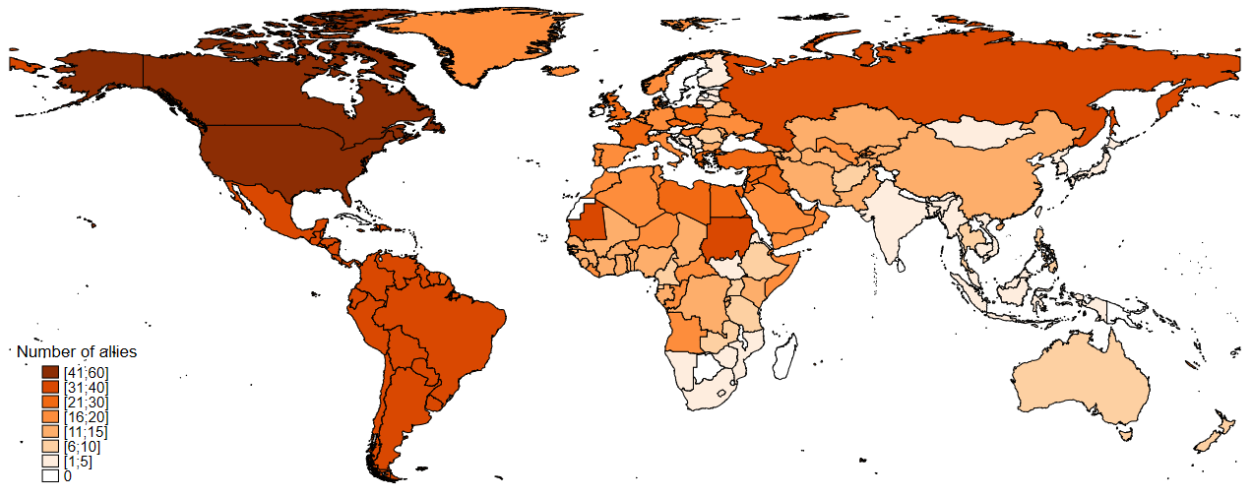
	Alliances	Defence pacts	Weak alliances
Countries treated	72	63	45
Countries never treated	12	21	38
Country-pairs treated	738	618	142
Country-pairs with a switch	362	244	142

*Note:* Author’s calculation. We count each exporter-importer observation as a country-pair. All alliances are symmetric. Ex: 72 countries of our sample have an active alliance with at least one partner for at least one year in our panel.

A simple density graph (cf. figure 1.3) displays a positive correlation between bilateral exports and military alliances. The distribution of country-pair exports with military alliances lies more to

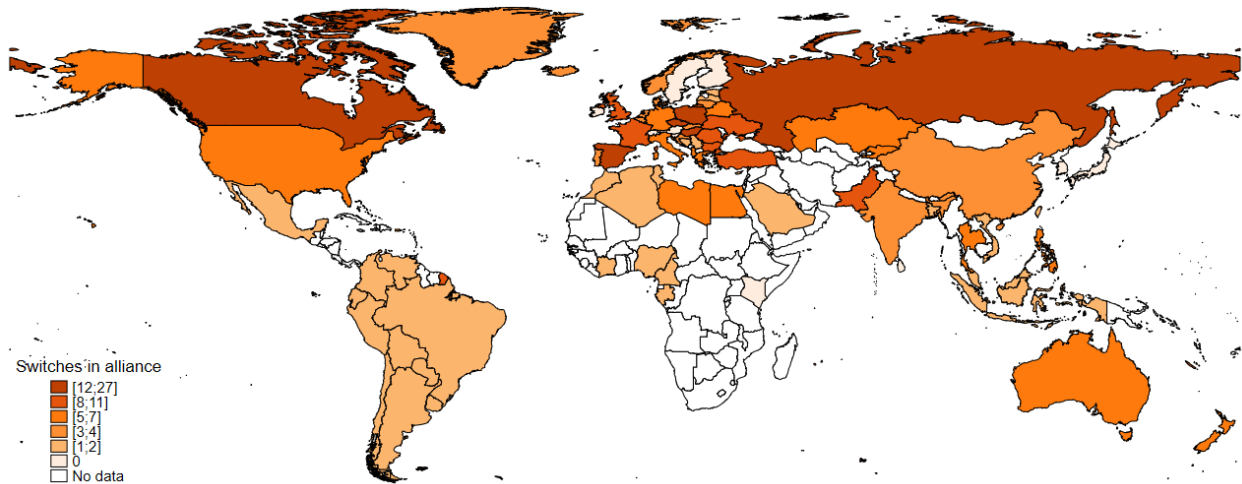
<sup>16</sup>United Nations Conference on Trade and Development Trade Analysis Information System.

Figure 1.1: Number of alliances by country, 1967-2012 (COW data)



*Note:* alliances are counted at country-pair level; 60 means that the country has been allied with 60 other countries from 1967 to 2012; white areas are where no alliance has been observed.

Figure 1.2: Sample switches in alliance by country

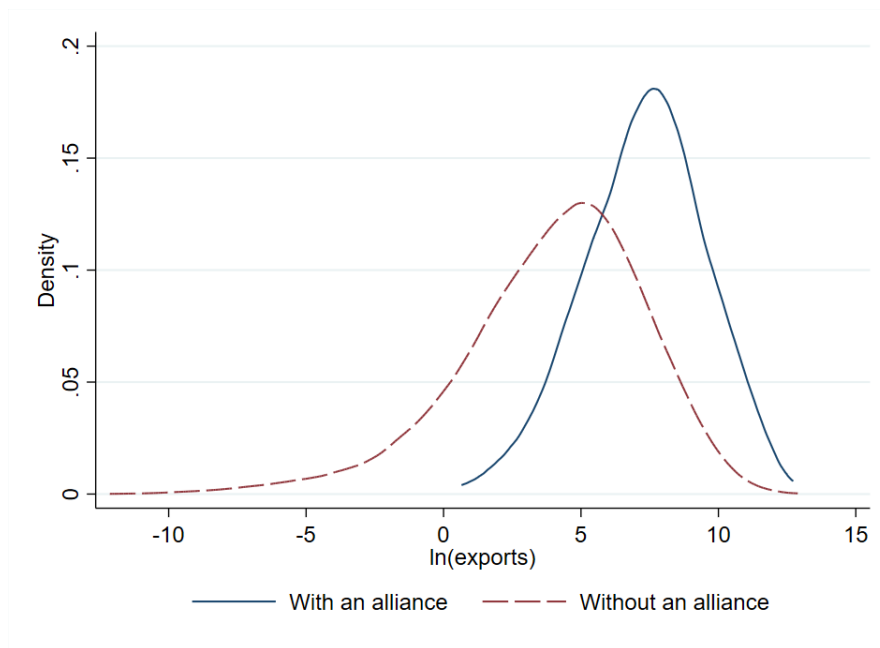


*Note:* Alliance switches are counted at country-pair level; 27 means that the country has signed or terminated a military alliance with 27 other countries from 1967 to 2012; white areas are countries excluded from our final sample.

the right than the distribution without, indicating a significantly higher average level of exports for country pairs with a military alliance than without.

As shown by Vicard (2012), the signature of RTAs is linked to the risk of interstate conflict. Yet, the correlation between alliance enforcement and exports does not seem to be very dependent on the existence of RTAs. Figure 1.4 presents a further two density graphs. Graph 1.4a, displays export values depending on whether the pair has an RTA, an alliance, both or no agreement. Graph 1.4b reproduces graph 1.3 but with export values conditional on RTAs (i.e. exports unexplained by RTAs). Both graphs present interesting evidence that, irrespective of the existence of an RTA, enforcement of an alliance is positively correlated with bilateral exports.

Figure 1.3: Export values and alliances



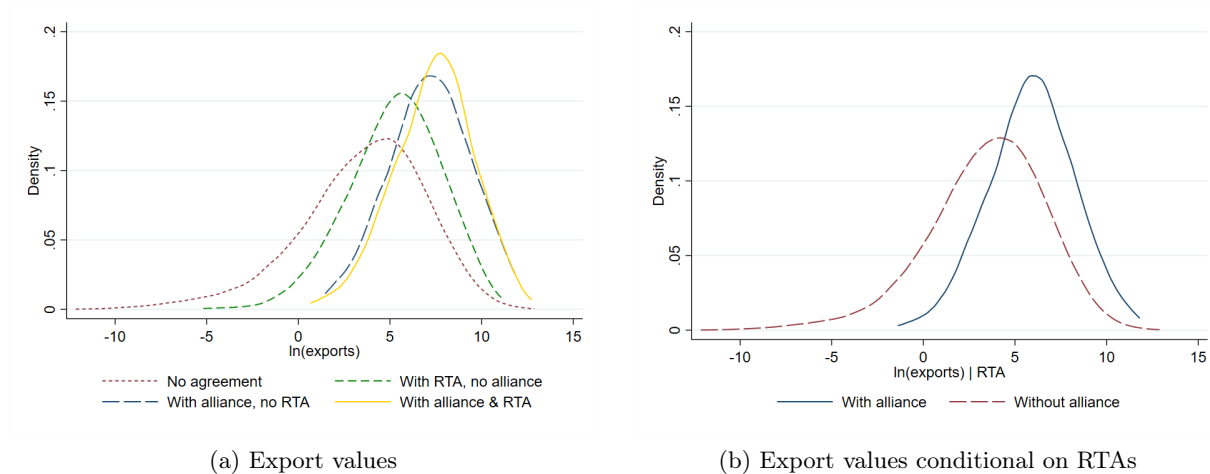
*Note:* The K-density graph compiles exporter-importer exports for the latest year of available data (2012). It displays the distribution of country-pair exports depending on the presence or the absence of an alliance.

### The role of alliances in insecurity

To pursue this descriptive discussion about the link between alliances, RTAs and trade, we regress in table 1.2 standard proxies of insecurity on  $ALL_{ijt}$  and  $RTA_{ijt}$ . These insecurity proxies are drawn directly from the literature presented in the introduction. First, we replicate the approach taken by Anderson and Marcouiller (2002) using an institutional proxy designed to capture the level of corruption. As in Anderson and Marcouiller (2002), institutional changes are observed at country-level and measured in log-ratios.<sup>17</sup> In addition, country i and j corruption log-ratios are interacted to obtain

<sup>17</sup>Formally:  $\ln\left(\frac{\text{corruption}_{it}+1}{E_{Wt}\text{corruption}_{it}+1}\right)$  and  $\ln\left(\frac{\text{corruption}_{jt}+1}{E_{Wt}\text{corruption}_{jt}+1}\right)$ , where  $W$  means the world.

Figure 1.4: Export values, alliances and RTAs



*Note:* K-density graphs compile exporter-importer exports for latest year of available data (2012). Graph (a) displays the distribution of country-pair exports depending on the presence or the absence of an alliance or a Regional Trade Agreement. Graph (b) displays the distribution of country-pair exports unexplained by Regional Trade Agreements depending on the presence or the absence of an alliance.

bilateral changes. Second, in keeping with [Rohner et al. \(2013\)](#); [Yu et al. \(2015\)](#), we derive insecurity costs from conflict signals, which are observed in terms of institutional law enforcement differences. Thus, in column (2), we use as a proxy for insecurity the log of the absolute-value difference in rule of law between countries  $i$  and  $j$ . Third, as in [Blomberg and Hess \(2006\)](#) we create a conflict dummy taking the value 1 if country  $i$  and  $j$  are both exposed to at least one conflict event at time  $t$  and 0 otherwise.<sup>18</sup> In this case, bilateral insecurity is measured using the extensive margin of country conflict exposure. The finding is the same in each column of table 1.2: the coefficient of  $ALL_{ijt}$  is statistically significant, negative and much larger than for  $RTA_{ijt}$  (see F-tests).<sup>19</sup> In other words, alliances are closely correlated with a lower level of bilateral insecurity, while RTAs are not.<sup>20</sup>

<sup>18</sup>Contrary to [Blomberg and Hess \(2006\)](#), the Conflict dummy used here is not restricted to non-state events. Thus, the variable also takes the value one in the case of violent events involving state forces.

<sup>19</sup>Note that we do not use the obvious War dummy measure as a proxy. Unlike [Glick and Taylor \(2010\)](#), our modern timeframe means that our panel contains very few interstate war observations. Moreover, the correlation between alliances (especially defence pacts) and war is very close to -1, since almost no country-pair with a switch in  $ALL_{ijt}$  was involved in an open war with each other. Therefore, we cannot regress a War dummy on alliances. Note here also that modern insecurity takes neither exclusively nor mainly the form of open interstate wars, but also exposure to (potential) violence from states or organised non-state actors for whatever reason (geo-strategic, economic, ideological, etc.).

<sup>20</sup>Given that, as seen from our theoretical discussion, insecurity can lead to market entry barriers imposed by states, it is not surprising to observe that RTAs are not completely uncorrelated with insecurity variables since these treaties (especially the deepest) can include some agreements on barriers that affect the fixed costs.



Table 1.2: Descriptive correlations: bilateral insecurity, alliances and RTAs

Dependent variable:	(1) Corruption	(2) Rule of law(diff.)	(3) Conflict dummy
Alliance	-0.278 <sup>a</sup> (0.050)	-0.066 <sup>a</sup> (0.007)	-0.088 <sup>a</sup> (0.013)
RTA	-0.003 (0.037)	-0.037 <sup>a</sup> (0.003)	-0.013 <sup>b</sup> (0.014)
Controls	yes	yes	yes
Country i x Year FE	no	yes	yes
Country j x Year FE	no	yes	yes
Dyadic FE	no	yes	yes
No. observ.	163,484	262,324	167,328
Alliance-RTA F-test	18.38	26.63	27.60

*Note:* The estimator is Ordinary Least Squares; Dependent variables are different proxies for bilateral insecurity as detailed in section 1.2. Robust standard errors clustered at country-pair level are in parentheses. Controls are  $\ln(GDP_{it})$ ,  $\ln(GDP_{jt})$ ,  $\ln(pop_{it})$ ,  $\ln(pop_{jt})$ ,  $\ln(distance_{ij})$ ,  $common.religion_{ij}$ ,  $common.language_{ij}$ ,  $colonial.past_{ij}$ ,  $contiguity_{ij}$ ; they are all captured by fixed effects in estimations (2) and (3). Difference in observation numbers are due to data availability restrictions. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively.

### 1.3 Identification strategy

Following our theoretical discussion, the relationship between military alliances and exports can be estimated using a structural gravity model. Accordingly, our baseline specification is as follows:

$$X_{ijt} = \exp(\beta_1 ALL_{ijt} + \beta_2 RTA_{ijt} + \lambda_{it} + \lambda_{jt} + \lambda_{ij}) * \epsilon_{ijt} \quad (1.3)$$

Our interest variable  $ALL_{ijt}$  is a dummy taking the value one if there is an alliance between country  $i$  and  $j$  at time  $t$  and zero otherwise.  $RTA_{ijt}$  is coded the same way as alliances, but for regional trade agreements. Alliances and RTAs may exist concurrently. Hence, we need to control for RTAs to capture any specific trade agreement effect between  $i$  and  $j$ .

$\lambda_{it}$  and  $\lambda_{jt}$  the exporter-year and importer-year fixed effects. They capture the country, year and country-year-specific variables such as economic size and multilateral resistance terms (Baier and Bergstrand, 2007; Feenstra, 2015; Redding and Venables, 2004).<sup>21</sup> We also include exporter-importer fixed effects ( $\lambda_{ij}$ ) to capture any omitted variables due to structural relations between countries such as distance, common language and colonial past. Military alliances active throughout the period are also captured. We hence estimate the within-effects of military alliances (i.e. country pair changes in status).<sup>22</sup>

<sup>21</sup>With this strategy, we focus on the bilateral effect of alliances (see the theoretical discussion in section 1.1). Any effect of alliances on trade through a reduction in multilateral insecurity costs are captured by the country-year fixed effects.

<sup>22</sup>We use a Poisson Pseudo Maximum Likelihood (PPML) estimator to retain a non-linear specification and address heteroscedasticity. In this way, we take into account zero trade observations and avoid the biases caused by a combination of log-linearisation and heteroscedasticity (Silva and Tenreyro, 2006). In addition, standard-errors are clustered at

Military alliances are expected to impact on all sectors by reducing insecurity costs. Yet, military alliances can also be associated with arms supply contracts. We therefore exclude the arms sector from the bilateral exports variable  $X_{ijt}$  to make sure that what is measured is a trade cost reduction and not a contract effect.

The reverse causality argument is unlikely to bias our baseline estimation. An alliance is not an economic treaty, but a long-lasting military pact with heavy political constraints. Therefore, to find a pair-specific export shock affecting the signature of a military alliance is a remote possibility. Nonetheless, we test alternative specifications and address residual endogeneity concerns in a further section.

### 1.3.1 Baseline Results

Table 1.3 reports the baseline results. The dependent variable is the exports from country  $i$  to country  $j$  in year  $t$ . The effect of military alliances on bilateral exports is positive and significant. Enforcing a military alliance increases bilateral exports by 60% on average. By contrast, the average effect of RTAs is 17%.<sup>23</sup> Translating the effects of alliances into tariff-equivalent variations under the standard trade elasticity calibration of  $\theta = 3.7$ <sup>24</sup> returns a tariff reduction of 12.8%. This equivalence appears reasonable given that we are dealing with a treatment assumed to have a sizeable impact on bilateral insecurity.<sup>25</sup> In subsequent sections, we are careful to test the robustness of this result and focus on understanding the heterogeneity and mechanism behind it.

### 1.3.2 Sensitivity analysis

This section proposes a battery of robustness checks for our baseline results.

*Intra-national trade flows.*— Data limitations are such that using intranational trade flows would significantly reduce our sample and limit the observation of military alliances’ within-variations (see section 1.2 on data for more details). Thus, the baseline estimation considers international trade flows only. However, this means that the alliance coefficient cannot include the potential effect of alliances in terms of diverting trade from intranational flows to international flows (Dai et al., 2014). Therefore, as a robustness check, we replicate the baseline estimation using a smaller panel containing within-country exporter-importer level.

<sup>23</sup>Recently, Larch and Yotov (2023) reviewed estimated effects of RTAs in the literature depending on estimation strategies. They show that in structural gravity with PPML estimator, RTAs’ coefficient should be around 0.14. With a coefficient of 0.16, our RTA estimate is in line with this general result.

<sup>24</sup>We directly estimate trade elasticity by including tariffs in the standard structural gravity estimation – see appendix table 1.G.13. The coefficient of  $\ln(1 + tariffs)_{ijt}$  is directly interpretable as the trade elasticity (Anderson et al., 2018; Head and Mayer, 2014)

<sup>25</sup>In the literature, the negative effect of bilateral insecurity on trade is estimated between -35% and -80% depending on the proxy (Martin et al., 2008; Glick and Taylor, 2010). This assumption is directly treated (and confirmed) in section 1.6

Table 1.3: Exports and military alliances

Estimator:	PPML
Dependent variable:	Bilateral exports
Variables	(1)
Alliance	0.473 <sup>a</sup> (0.106)
RTA	0.161 <sup>a</sup> (0.033)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country  $i$  to country  $j$  at time  $t$  in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

trade flow data. Results are displayed in 1.G.2 column 1.

*Tariffs*– We investigate the robustness of our results to the inclusion of tariffs. The tariff variable is defined as the log of the ad valorem exporter-importer-time average tariff rate plus one. The results of this sensitivity test are reported in table 1.G.2 column 2 to 4. The tariff coefficient is negative and significant. This additional control variable does not reveal any omitted variable bias. The lower coefficient of the alliance variable is merely the outcome of the considerable reduction in the timeframe and the country-pairs sample for our panel dictated by tariff data availability. To overcome this lack of country-pair-within variation due to the sample size, we include within-country trade flows – see column 5. This strongly increases alliances’ coefficient as well as its precision. In addition, controlling for common GATT membership has no effect.<sup>26</sup>

*RTA depth*– In the baseline estimation, dummy variable  $RTA_{ijt}$  for the presence of an RTA between exporter and importer. This ensures that the coefficient of  $ALL_{ijt}$  is not biased by concomitant variations in RTAs and alliances. Yet, this method does not control for changes in RTA depth, which may be correlated with conflict related variables (Vicard, 2012). To address this point, we proxy RTA depth by the number of provisions in each agreement. We then introduce this new variable into our baseline estimation. We drop country-pair-year observations where an RTA is observed, but not its depth. Results are displayed in table 1.G.2 columns 6-7. We keep the  $RTA_{ijt}$  dummy in column one. Its coefficient is not empirically interpretable since it corresponds to a fictive empty RTA (with no

<sup>26</sup>We introduced a dummy variable taking value one if both country  $i$  and  $j$  are members of the GATT. The alliance coefficient and standard error are unaffected by this additional control.

provision). In the second column,  $RTA_{ijt}$  is dropped. In both cases, the coefficient of  $ALL_{ijt}$  is barely affected.

*Distance and economic development.*— Geographic distance and long-term differences in economic development are already captured by the country-pair fixed effect. Yet, both bilateral exports and signatures of alliance treaties could be affected by regionalisation or globalisation, i.e. by variations in transport costs and differences in economic development over time. Therefore, we build  $\ln(\text{distance}_{ij}) * \text{year}_t$  as an interaction variable between the distance and the year <sup>27</sup>, and  $\ln(|\frac{GDP_{it}}{Pop_{it}} - \frac{GDP_{jt}}{Pop_{jt}}|)$  as the log of the difference in per capita GDP in absolute value between exporter and importer. We include these variables in the baseline estimation. Results are reported in table 1.G.2 column 8. Their coefficients are weak and not significant, while the military alliance coefficient remains unchanged.

*Extended Panel.*— For quality reasons, our trade data limit our panel’s sample of country-pairs and time-frame. We test the sensitivity of our results to alternative trade data allowing for a larger panel. We use the Correlate of War trade database (Keshk, 2017; Barbieri et al., 2009) to extend our panel from 46 to 143 years, and from 6,972 to 36,737 country pairs. Yet there are costs involved. First, we do not observe all country-pairs over the entire period. Second, data are not subject to the same harmonisation and verification process as CHELEM. This undermines quality, especially for developing countries. Third, we cannot exclude the arms sector. Results are reported in table 1.G.2 column 9 and 10.<sup>28</sup> We still observe a clear positive and significant effect of alliances on trade. Nonetheless, the data limitation and the panel change mean that we cannot directly compare with the baseline. After 2012, our panel is limited by alliances data. We proceed to an additional sensitivity test by extending our panel until 2020. We complete our sample using Alliance Treaty Obligations and Provisions (ATOP) data (Leeds et al., 2002).<sup>29</sup> ATOP has a larger time-coverage than the Correlate of War alliances data. Yet, ATOP is less precise and does not provide clear information about the content of treaties. Results are displayed in 1.G.2 column 11. The alliance coefficient is poorly affected by this additional panel extension.

*Asymptotic bias.*— Table 1.G.3 presents the corrected FE-PPML estimation developed by Weidner and Zylkin (2021). The military alliance coefficient is affected by a small negative bias of the order of -0.01, while the associated standard error is slightly underestimated. Once the correction is applied, the alliance coefficient remains highly significant and very similar to the baseline.

*Negative weights.*— The baseline strategy is similar to a fixed-effects difference-in-differences es-

<sup>27</sup>For ease of coefficient interpretation, the year variable is equal to the year minus 1966.

<sup>28</sup>In column 9, we do not control for RTAs to use the largest panel. Controlling for RTAs in column 10 makes start the panel in 1948 instead of 1870.

<sup>29</sup>ATOP data stop in 2018. Yet, between 2019 and 2020, no alliance treaty has been reported by the United Nations. We assume, therefore, that the alliances active in 2019 and 2020 are the same as in 2018.

mination. Yet, the effects of military alliances could be dynamic.<sup>30</sup> In this case, a different effect would be found depending on the duration of the treaty, where the baseline coefficient would be the average. Moreover, the effects of military alliances on trade could also be heterogeneous across time and country pairs.<sup>31</sup> Therefore, our results could suffer from negative-weight biases (De Chaisemartin and d’Haultfoeuille, 2020). Thus, to address these econometric considerations, we regress equation (1.4) using the De Chaisemartin and d’Haultfoeuille (2024) estimator. The estimation is detailed in appendix section 1.D. It concludes that signing an alliance increases bilateral exports by 40% at time  $t$  (the date of signature). The effect then gradually grows over time to attain 98% in  $t + 5$ . Since the baseline results are the average of the effects of military alliances over time, the estimated effect of 60% is consistent with the observed dynamic and robust to the negative weights bias. De Chaisemartin and d’Haultfoeuille (2024)’s estimator is linear, which in a gravity setting may bias results (Silva and Tenreiro, 2006). Following Nagengast and Yotov (2023); Wooldridge (2023)’s method, which allows for ppml estimator, we confirm our results.

### 1.3.3 The heterogeneous effect of alliances

#### Heterogeneity of alliance treaties

There can be two types of alliances: *the weak alliances*<sup>32</sup> and *the defence pacts*. This distinction is based on the particularity of defence pact security cooperation policies (Gibler, 2008). In a nutshell, defence pacts are the only alliances designed to protect members from (common) external threats (see appendix section 1.A for more details). We do not consider this distinction in the baseline. Yet, because of this fundamental treaty difference, we might expect defence pacts to have more of an effect on trade than weak alliances. Long (2003) provides a test of this difference using a non-structural gravity approach, concluding that only defence pacts are positively correlated with trade. In this section, we investigate whether the difference in alliance treaty categories produces heterogeneous effects, estimating within-effects using a structural gravity model.

We decompose the alliance variable into: i) a dummy taking the value one if there is an alliance with a defence pact between  $i$  and  $j$  at time  $t$ , and ii) another dummy taking the value one if there is an alliance without a defence pact between  $i$  and  $j$  at time  $t$ . Then, we replicate our baseline estimation. The results are reported in appendix table 1.G.4 column 1. Defence pacts increase bilateral exports by 100%, while the weak alliance coefficient is not statistically significant. This shows that a high level of military cooperation policies is required for alliances to affect trade. Moreover, the fact that defence pacts account for the majority of alliances explains the intensity of our baseline result – the

<sup>30</sup>We properly discuss the dynamic effect of alliances in section 1.5.

<sup>31</sup>We extensively discuss the heterogeneous effect of alliances in the following section 1.3.3.

<sup>32</sup>Non-aggression pacts, neutrality pacts and ententes.

latter is driven by the effect of very deep agreements. Conclusions are robust to the Cold War and the particularities of international relations during this period. A full discussion about the Cold War is provided in appendix section 1.E. In addition, we test whether our results are driven by the largest defence pact in our sample – NATO. We create a specific dummy variable for the treaty and retain the defence pact dummy variable for the others.<sup>33</sup> The estimation results presented in table 1.G.4 column 2 confirm our conclusions.<sup>34</sup>

### Bilateral elasticities

In line with standard gravity theory, trade cost elasticities – including insecurity cost’s one – are constant in our model (Head and Mayer, 2014).<sup>35</sup> Yet, recent literature has shown that the trade elasticity is decreasing with the value of bilateral trade (Carrère et al., 2020; Chen and Novy, 2022). As developed in section 1.1, military alliances are expected to affect trade by reducing trade costs’ sensitive to insecurity. Therefore, this section investigates whether the effects of military alliances on bilateral exports are concerned by bilateral trade elasticities.

In keeping with Carrère et al. (2020), we use the Method of Moments-Quantile Regression estimation (Machado and Silva, 2019):

$$\tilde{X}_{q,ijt} = \beta_1 ALL_{q,ijt} + \beta_2 RTA_{q,ijt} + BZ_{q,ijt} + \lambda_{q,it} + \lambda_{q,jt} + \epsilon_{q,ijt} \quad (1.4)$$

$\tilde{X}_{q,ijt}$  is the inverse hyperbolic sine transformation of  $X_{q,ijt}$ .  $\lambda_{q,it}$  and  $\lambda_{q,jt}$  are exporter-year and importer-year fixed effects, while  $q$  is a defined quantile and  $Z$  a vector of bilateral variables: distance, interaction between distance and year, colonial past, common language, common currency, common religion and territorial contiguity. Given the sub-divisions performed by the quantile estimator, introducing dyadic-FE would entirely capture the effect of our variable of interest. The estimator is linear. So, the dependent variable is the natural log of exports augmented by one. We regress the equation for each percentile. The obtained  $ALL_{q,ijt}$  coefficients, with bootstrapped 95% confidence intervals, are graphically displayed in figure 1.5. Similarly to the conclusions drawn by Carrère et al. (2020) with respect to distance, the alliance coefficients are decreasing in the value of trade. We plot the RTA coefficients in appendix graph 1.G.2. They also show a decrease in coefficients with the size of exports.<sup>36</sup>

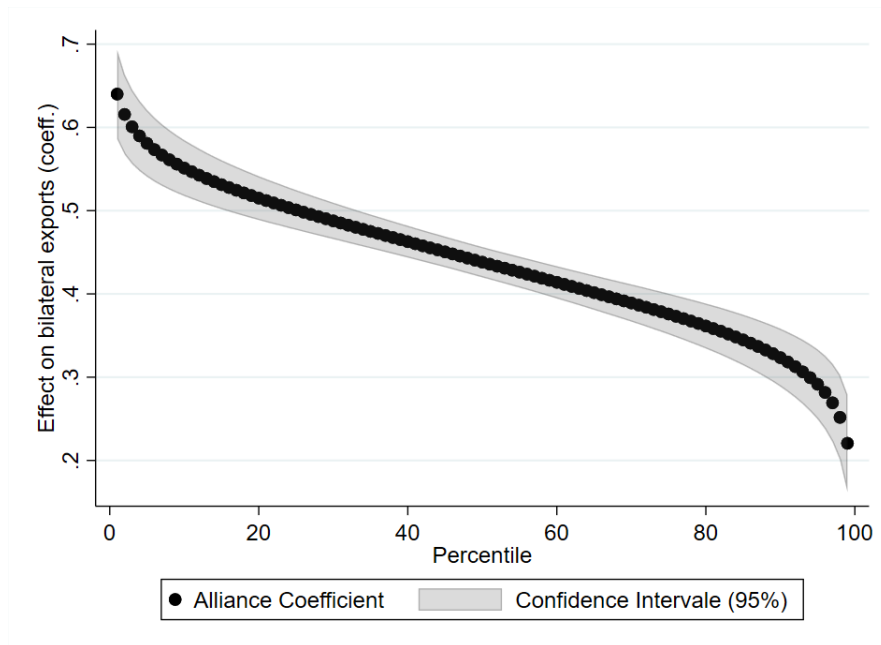
<sup>33</sup>Thus, in this estimation, the defence pact variable takes the value 0 if  $NATO = 1$ .

<sup>34</sup>In keeping with the definition of military alliances given by Gibler (2008), some recent NATO members (post-2003) are considered to belong to the international organisation, but not the alliance. We test the validity of our estimation including all NATO members in the alliance system. Results are barely affected.

<sup>35</sup>In equation 1.3, trade elasticities are determined by  $\theta$ , the Pareto shape parameter of the productivity distribution which assumed to be constant across country-pairs.

<sup>36</sup>Another way to test the presence of bilateral trade elasticities is to change the weight of observations in our baseline. Using import share as the dependent variable, we increase the weight of observations with small export flows (Sotelo, 2019). We report the results in appendix table 1.G.2 column 10. In line with our findings, we observe higher coefficients

Figure 1.5: Quantile estimates of alliances



Two main conclusions are drawn from these results. First, the effect of military alliances on trade depends negatively on the value of bilateral flows. This implies that small country-pairs benefit the most from the signature of a military alliance. Second, military alliance coefficients behave in the same way as variables that clearly affect trade through trade costs. This points to an alliance effect through a trade cost reduction mechanism. In appendix section 1.B.3, we discuss the consistency of the bilateral trade elasticities with our theoretical framework and provide an extension of our gravity model.

## 1.4 Endogeneity

This section takes the identification of the causal effects of military alliances on trade a step further. Indeed, we still need to address potential reverse causality (a trade shock affecting the alliance signature decision) and country-pair-time varying omitted variables (such as a bilateral geopolitical shock). For this purpose, we develop an instrumental variable approach based on common outside alliances. The intuition behind this is the domino-like spread of international agreements, as tested for RTAs or regionalism by Baldwin and Jaimovich (2012). Common allies present similar military interests and the inability of one country in the pair to use its alliances in a conflict against the other. Therefore,

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for both alliances and RTAs. Furthermore, adding in our baseline an interaction between  $ALL_{ijt}$  and country-pair specific weights  $(\frac{\sum_{k \neq j} X_{ik,t-1}}{\sum_k X_{ik,t-1}} * \frac{\sum_{k \neq i} X_{kj,t-1}}{\sum_k X_{kj,t-1}})$  capturing to what extent each partner is relatively important for the other in its total trade, we also observe a negative correlation between the alliance effect and the relative importance of  $X_{ij}$ .

if countries  $i$  and  $j$  have many allies in common, it is highly probable they will develop an alliance together, while in partial equilibrium (i.e. controlling for aggregated prices, output and expenditures) the change in bilateral insecurity costs with external partners should not directly affect the trade inside the pair. We want to avoid capturing the inverse relation: an increase of the common outside alliances because of the signature of an alliance between countries  $i$  and  $j$ . Therefore, we ignore variations in common outside alliances while countries  $i$  and  $j$  are allied.<sup>37</sup> Additionally, we exclude also from the sum the alliance between country-pair  $ij$  – the country-pair of interest. Thus, we can write the instrumental variable as follows:

$$IV_{ijt} = \begin{cases} \sum_{k \neq i, j; t_{all}} (ALL_{ikt_{all}} * ALL_{kjt_{all}}), & \text{if } \exists t_{all}, t > t_{all} \ \& \ ALL_{ijt} = 1 \\ \sum_{k \neq i, j; t} (ALL_{ikt} * ALL_{kjt}), & \text{otherwise} \end{cases} \quad (1.5)$$

We use an OLS/PPML two-stage approach to keep the IV strategy comparable with the baseline and prevent log-linearisation under heteroscedasticity from biasing our estimates (Silva and Tenreyro, 2006). In the first stage, we use OLS to estimate the effect of the IV on the probability of signing an alliance. The predicted probability is used to compute the instrumented alliances. Then, we estimate with a PPML the effect of the instrumented alliances on bilateral exports. In the second stage, clustered standard-errors are bootstrapped.<sup>38</sup> Results are reported in table 1.4. The (instrumented) military alliance coefficient is strongly positive and significant.<sup>39</sup> Therefore, the IV estimation confirms the causal interpretation of the effects of military alliances on trade.<sup>40</sup>

Using a non-linear second-stage estimator may induce consistency issues. Lin and Wooldridge (2019) recommend a control-function approach to address this. Instead of using alliances' predicted probabilities, we include first-stage residuals as control. Results, reported in appendix table 1.G.6, confirm the OLS/PPML two-stage conclusions. In addition, the control-function approach allows testing the baseline's sensitivity to omitted variable concerns. The coefficient of the first-stage residuals is directly interpretable as the omitted variable bias addressed by the IV. Interestingly, focusing on

<sup>37</sup>Formally, as long as countries  $i$  and  $j$  are allies, we set the value of the sum of their common outside alliances at the year of signature  $t_{all}$

<sup>38</sup>Bootstrapping not only the second stage but the whole process does not affect results.

<sup>39</sup>We observe a higher coefficient compared with the baseline results. This is due to alliance selection induced by the IV. By targeting alliances included in an international network, we mechanically select defence pacts, which tends to increase coefficient (see section 1.3.3). Reduction in second-stage standard errors results simply from bootstrapping.

<sup>40</sup>Supplementary sensitivity tests are performed on the IV results. Results are robust to the inclusion of an interaction variable between distance (in log) and year, which controls for the globalisation dynamic. To avoid potential biases due to RTA endogeneity, we also run an estimation in which we instrument them by common outside RTAs ( $\sum_{k \neq i, j; t} RTA_{ikt} * RTA_{kjt}$ ). This does not affect the instrumented alliance coefficient. We provide a 2SLS estimation in appendix table 1.G.5. The alliance coefficient is still positive and significant. Yet, the OLS and 2SLS gravity estimations induce biased coefficients and standard errors – due to log-linearisation under heteroscedasticity (Silva and Tenreyro, 2006) – and observation weights different to PPML (Sotelo, 2019).



defence pacts (column 2), the residuals' coefficient is significantly smaller, suggesting that defence pacts are less concerned by omitted variable issues than weak alliances. <sup>41</sup>

Table 1.4: Alliances and bilateral exports, IV

Dependent variable:	exports
Second stage	(1)
Estimator:	PPML
Instrument variable:	Common out. alliances
Alliance	0.655 <sup>a</sup> (0.026)
RTA	0.159 <sup>a</sup> (0.011)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666
First-stage	
Estimator:	OLS
Instrumented variable:	Alliance
Common out. alliances	0.056 <sup>a</sup> (0.007)
RTA	0.011 (0.008)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.712
KPW F-stat	67
KPW LM-stat	11

*Note:* OLS, Ordinary Least Squares; FE, Fixed effects. Dependent variable is exports from country  $i$  to country  $j$  at time  $t$  in millions of current dollars. Standard errors clustered at the exporter and importer levels are in parentheses. Second-stage standard errors are bootstrapped. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively.

### 1.4.1 IV validity

The IV approach is based on the assumption of non-violation of the exclusion restriction. The common outside alliances can impact country  $i$ 's exports to  $j$  only through enforcement of an alliance. If, as discussed the theoretical section 1.1, in partial equilibrium alliances affect trade only through a reduction in bilateral insecurity costs, our IV should respect this restriction. Yet, in a logic of geopolitical alignment on international scene, the signature of an alliance may lead to a variation in trade costs with non-signatory partners. To qualitatively test it, we create two dummies:  $ALL_{i,-j,t}^{out}$  and  $ALL_{-i,j,t}^{out}$ .

<sup>41</sup>Furthermore, coefficients of instrumented alliances and defence pacts are very similar, which supports that defence pacts drive the estimated causal effect on trade.

$ALL_{i,-j,t}^{out}$  takes value 1 if the exporter has signed an alliance with any country other than  $j$  (i.e. an outsider) and zero otherwise. Similarly,  $ALL_{-i,j,t}^{out}$  takes value 1 if the importer has signed an alliance with any country other than  $i$ . Both  $ALL_{i,-j,t}^{out}$  and  $ALL_{-i,j,t}^{out}$  can be estimated in the presence of exporter-year, importer-year and country-pair fixed effects since they are country-pair-year specific (Dai et al., 2014). Both dummies are introduced in equation (1.4) estimated by a PPML. We report the results in appendix table 1.G.7.<sup>42</sup>  $ALL_{i,-j,t}^{out}$ 's and  $ALL_{-i,j,t}^{out}$ 's coefficients are not significantly different from zero. So, on average, signing an alliance with an outsider does not affect exports from country  $i$  to  $j$ .<sup>43</sup> This result points to the validity of the instrument's exogeneity assumption.

Furthermore, we perform a *plausible exogeneity test* proposed by Conley et al. (2012). The approach aims at supporting the validity of our results even in the presence of a reasonable deviation from perfect alignment with the exclusion restriction assumption. We relax the assumption by allowing for  $\nu$ , the correlation between the instrumental variable (common outside alliances) and errors, to deviate from zero. We then test whether the estimate of the instrumented variable ( $ALL_{ijt}$ ) is robust to such a deviation from the exclusion restriction.

The union of confidence interval method calls first for setting the minimum (or maximum) value  $\nu$  can take. To approximate this, we regress bilateral exports on the endogenous variable ( $ALL_{ijt}$ ) and the instrumental variable (common outside alliances) with our standard set of fixed effects and controls. The coefficient associated with the IV represents an approximation of the degree of deviation from the exclusion restriction (Kippersluis and Rietveld, 2018). We obtain a small coefficient (-0.030). We then plug this degree of deviation into the plausible exogeneity test. In this way, we obtain the estimation of the (instrumented) alliance coefficient's upper and lower bounds under the relaxation of the exclusion restriction (see table 1.5).<sup>44</sup> The interval [0.388 ; 1.089] does not contain zero. Thus, we can safely argue that military alliances have an unambiguous positive causal effect on bilateral exports.

## 1.5 The dynamic effect of alliances

So far, we have focused on the within-average effect of alliances. However, alliances may require time to be fully effective and increase bilateral trade dynamically. Applying a simple event study approach on our baseline does not permit to identify the dynamic effect of alliances in the absence of pre-trend (see

<sup>42</sup>We run the same estimation in column two, but decompose the  $ALL_{ijt}$  dummies into defence pacts and weak alliances. This does not affect the results.

<sup>43</sup>Our time frame include the Cold-War. During this period because strong geopolitical *blocks* of countries existed, signing an alliance with a country of on block may reduce trade with members of the other one. We further discuss this issue in the appendix section 1.E. We show that in this specific context such a reduction in trade applies, but also that it the Differenced Average Treatment on the Treated approach we develop in a further section (1.5) perfectly addresses this issue.

<sup>44</sup>The Conley et al. (2012) estimator is based on the 2SLS estimator. We therefore have to log-linearise the gravity equation. We use the inverse hyperbolic sine transformation of exports to capture zero values. As discussed previously, log-linearisation under heteroscedasticity could bias the alliance coefficient. Nevertheless, estimated bounds are consistent with the OLS/PPML two-stage estimation which addresses this bias.

Table 1.5: Plausible exogeneity test

Dep var: exports	<i>Union of Confidence Interval estimations</i>		
Instrumented var.	$\nu$	Min	Max
	-0.030	95% CI	95% CI
Alliance		0.408	1.106

*Note:* UCI based on the IV's  $\nu$  coefficient from a regression of exports on interest variables and the IV

graph 1.G.3 in appendix). Therefore, we develop in this section a Differenced Average Treatment on the Treated (DATT) approach based on Couch and Placzek (2010). In the light of the results in section 1.3.3, we choose to increase the precision of the analysis by focusing on defence pacts. Therefore, we decompose variable  $ALL_{ijt}$  into *defence pacts* and *weak alliances*. Then, we estimate the defence pact DATT, using weak alliances as a control variable.

The observation of a pre-trend in the simple event-study suggests the presence of a selection bias. To address this issue, we perform dynamic propensity score matching on a staggered sample.<sup>45</sup> We define the propensity score  $p(x_{ijt})_t$  as the likelihood of signing a defence pact conditional on a set of standard observable gravity variables. The propensity score is estimated for each separate year from 1967 to 2012. In this way, depending on the year, the variables are allowed to affect  $p(x_{ijt})_t$  differently. Each treated country-pair observation is matched on  $p(x_{ijt})_t$  with an untreated country-pair. A country-pair with no defence pact, but having had one in the past is never used as a control observation. We obtain a set of country-pair-year observations matched on  $p(x_{ijt})_t$ .<sup>46</sup>

In keeping with Couch and Placzek (2010), we estimate the equation:

$$X_{ijt} = \exp\left(\sum_{k>k'} \delta^k D_{ijt}^k + \beta_2 \text{Weak.ALL}_{ijt} + \beta_3 \text{RTA}_{ijt} + \lambda_{it} + \lambda_{jt} + \lambda_{ij}\right) * \epsilon_{ijt} \quad (1.6)$$

$k'$  must be a year or period prior to the  $k$  of signature of a defence pact between countries  $i$  and  $j$ .  $D_{ijt}^k$  is a dummy variable equal to 1 if  $t$  is the  $k^{\text{th}}$  year after (or the  $k^{\text{th}}$  year before if - $k$ ) the signature of a defence pact between countries  $i$  and  $j$ .  $\text{Weak.ALL}_{ijt}$  and  $\text{RTA}_{ijt}$  are dummy variables controlling respectively for the presence of a weak alliance and an RTA between countries  $i$  and  $j$  at time  $t$ .  $\lambda_{it}$ ,  $\lambda_{jt}$  and  $\lambda_{ij}$  are respectively exporter-year, importer-year and exporter-importer fixed-effects. We use a PPML estimator.

The DATT compares the difference in bilateral exports between years  $k$  and  $k'$  for a country-pair signing a defence pact during our period, indicated by  $D_{ij} = 1$ , to the difference in bilateral exports

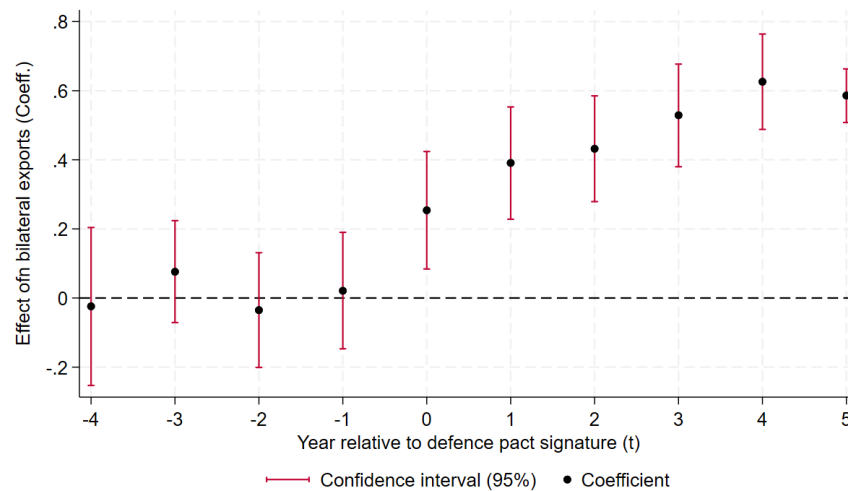
<sup>45</sup>This implies that previously treated country-pairs can never be part of the control group.

<sup>46</sup>See the appendix tables 1.G.8 and 1.G.9 for more details on dynamic propensity score matching. We evaluate the extent to which the matched treated and control groups are similar. We consider the standardised difference in means ( $B$ ) and the variance ratio ( $R$ ). To conclude that the groups are comparable,  $B$  must be inferior to 0.25, and  $R$  between 0.5 and 2 (Rubin, 2001; Stuart and Rubin, 2008). After matching, we obtain  $B = 24.7\% < 25\%$  and  $0.5 < R = 1.7 < 2$ . Thus, in the set of matched observations, the control and treated groups are well balanced.

between years  $k$  and  $k'$  for a country-pair with no defence pact, indicated by  $D_{ij} = 0$ . The expected difference between the year-to-year difference in exports is estimated for the set of treated country-pairs relative to the matched set of non-treated pairs (Couch and Placzek, 2010).

The results of the DATT estimation are graphically represented in figure 1.6. More details are provided in appendix table 1.G.10.  $t_{def}$  is the date of the defence pact's signature. We choose  $k' \leq t_{def} - 5$ . So, all  $k$  are estimated in comparison to  $k' \leq t_{def} - 5$  (the natural trend). As soon as the treaty is signed, the defence pact has a positive and significant effect on exports from  $i$  to  $j$ . In the following years, the effect keeps growing to stabilize in  $t_{def} + 5$ . Given that  $t_{def}$  the date of signature and not of enforcement, this period can be interpreted as the time of adjustment required for the defence pact to become effective and fully operational. For  $k \geq t_{def} + 5$  (the treated trend), the average estimated effect confirms our previous results: following their signature, defence pacts increase bilateral trade by 79%.<sup>47</sup> Moreover, for any  $k < t_{def}$ , we obtain a weak and insignificant coefficient. Therefore, the measured effect for any  $k \geq t_{def}$  is independent of any pre-trend.<sup>48</sup> Thus, in the light of our results, we can safely conclude that the defence pacts have an unambiguous gradual, positive and causal effect on trade.

Figure 1.6: Dynamic effect of defence pacts on bilateral exports (DATT)



*Note:* Figure plots results of the DATT estimation of the defence pacts' effect on bilateral trade. Every time  $k$  is compared to  $k' \leq t_{def} - 5$ . The last period includes  $k \geq t_{def} + 5$ . Estimator is PPML. Standard errors are clustered at the country-pair-year level.

<sup>47</sup>Since the data set is staggered, the DATT estimates only the entry effect (switches from 0 to 1). In previous estimation strategies, both entry and exit are considered.

<sup>48</sup>DATT results are also robust to the cold war. See the full discussion on the cold war in section 1.E.

## 1.6 From alliances to trade: the mechanism

This section analyses the mechanism behind the positive effect of alliances on trade.<sup>49</sup> As discussed in the theoretical section 1.1, alliances should affect trade by reducing the international insecurity between members. The first source of insecurity these treaties address is the risk of conflict between signatories.<sup>50</sup> Yet, several pieces of evidence support that this channel does not apply. First, in section 1.3.3, we show that the alliances focusing on the reduction on reducing the probability of conflict between signatories (i.e. the weak alliances) have no impact on bilateral trade. Second, in data, interstate conflicts are extremely rare in our panel.<sup>51</sup> Therefore, we do not observe that interstate conflicts are the counterparts of the absence of an alliance. Extending the risk of conflict in the country pair to geopolitical tensions, we estimate the effect of defence pacts on the common United Nations votes of the country-pair.<sup>52</sup> Results are reported in the appendix table 1.G.11. Defence pacts' effect is extremely low (+2%) while the weak alliances have no significant impact.<sup>53</sup> This supports that alliances do not increase trade by reducing the probability of conflict between signatories.

Nevertheless, defence pacts' content is not limited to the reduction of interstate conflicts between members. These alliances are designed to enforce deeper security cooperation aimed at addressing any source of insecurity that concerns both signatories. Indeed, as we can observe in the data, conflict events are, in the extremely large majority, related to fights against non-state organized actors (terrorism, civil wars, piracy, armed factions).<sup>54</sup> Taking into account the complexity of conflicts, we therefore use a broader measure of bilateral insecurity:

$$INS_{ijt} = \sum(\text{conflict\_event}_{it}) * \sum(\text{conflict\_event}_{jt}) \quad (1.7)$$

the interacted country-time sums of conflict events between  $i$  and  $j$ .<sup>55</sup> This measure has evident advantages over the proxies presented in section 1.2.2. First, it presents a good country-pair-time

<sup>49</sup>Most previous empirical papers focusing on the link between military alliances and trade consider that alliances affect exports because signatories are more inclined to reduce tariffs or sign RTAs with their allies (Long, 2003; Long and Leeds, 2006; Mansfield and Bronson, 1997). Yet, our baseline estimation includes RTAs as a control variable, while our analysis shows that our baseline results are barely sensitive to the inclusion of tariffs and RTA depth (see section 1.3). Therefore, this mechanism is excluded.

<sup>50</sup>See treaties description in appendix section 1.A.

<sup>51</sup>Using the Military Interstate Dispute database from the Correlate of War project (Palmer et al., 2022), we find in our sample only 86 observations (exporter-importer-time) with an active war. Following (Martin et al., 2008) by including lower magnitude conflicts (starting at simple display of force), we find only 1578 exporter-importer-time observations with an active interstate conflict – 436 happening while an alliance is active. Performing a linear probability estimation with our full set of fixed effects, we do not observe any significant effect of alliances on the probability of interstate conflict.

<sup>52</sup>For this purpose we use the United Nations General Assembly Voting Data from Voeten (2013).

<sup>53</sup>We estimate the effect of alliances on common UN votes at the country-pair-time level. The dependant variable being a count we use a PPML estimator. Since we want to capture the degree of geopolitical convergence the dependant variable includes only expressed votes while we control for abstention. In column (2) we use our instrumental variable. This does not affect our conclusions.

<sup>54</sup>Using UCDDP data – see section 1.2 – between 1989 and 2012 we observe 120.029 conflict events.

<sup>55</sup>Military cooperation events capture a reduction in insecurity rather than an increase. They therefore need to be excluded for an accurate measurement of bilateral insecurity.

variation. Second, it targets insecurity by directly focusing on exposure to violence. Third, it captures the extensive margin of bilateral insecurity (if countries  $i$  and  $j$  are exposed to insecurity) and the intensive margin (to what extent the country-pair is exposed).

We use two-stage estimations to look at the effect of defence pacts on exports by means of the variation of  $INS_{ijt}$  they enforce. The set of fixed effects is the same as in the baseline. This implies that the independent level of insecurity of country  $j$  or in country  $j$  at time  $t$  are controlled which ensures that we capture only the country-pair specific variation of  $INS_{ijt}$ . For interpretation reasons, we use the inverse hyperbolic sine transformation of  $INS_{ijt}$ .<sup>56</sup> Due to data limitations with respect to conflict events, we use a sub-sample starting in 1989. The results are displayed in table 1.6.<sup>57</sup> Column 1 presents the defence pact dummy variable. To further address possible endogeneity, we replace it in column 2 with common outside defence pacts.<sup>58</sup>

In the first-stage estimations, we observe a strong negative effect of defence pacts on bilateral insecurity. In column 1, we estimate that the enforcement of a defence pact reduces bilateral insecurity by 69%. Then, in the second-stage results, we observe in both specifications that the reduction in bilateral insecurity enforced by defence pacts (or common outside defence pacts) has a significant positive effect.<sup>59</sup> Thus, we estimate that, by reducing bilateral insecurity, defence pacts raise exports by 49.5%. In this sub-sample, this is equivalent to almost 100% of the total defence pact effect.

Alternatively, we test the insecurity mechanism estimating the effect of defence pacts on exports through the variation of  $\sum coop.mil.ev.ijt$ , the sum of the country-pair's cooperative military events. Contrary to  $INS_{ijt}$ , this does not directly target insecurity, but the enforcement of the military cooperation that reduces insecurity. Yet, we identify a purely bilateral variation. We report the two-stage estimations in appendix table 1.G.12.<sup>60</sup> Results strongly support both the validity and prevalence of the insecurity mechanism: by signing defence pacts, members sharply increase their effective security cooperation which ensures an important reduction in country-pairs insecurity and leads to an increase in trade between signatories.

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<sup>56</sup>We test different functional forms. This does not affect the robustness of our results. In addition, given that on average  $INS_{ijt} = 14622$ , we are not exposed to approximation errors using the inverse hyperbolic sine transformation (Bellemare and Wichman, 2020).

<sup>57</sup>As in section 1.4, we use a two-stage OLS/PPML estimator to prevent log-linearisation under heteroscedasticity from biasing our results (Lin and Wooldridge, 2019; Silva and Tenreyro, 2006). The control-function approach provides similar results.

<sup>58</sup>We construct this variable in the same way as the IV in section 1.4 but restricting it to defence pacts.

<sup>59</sup>The bilateral insecurity coefficient is negative since it expresses the effect of a *rise* in insecurity.

<sup>60</sup> $\sum coop.mil.ev.ijt$  measures the opposite of  $INS_{ijt}$ . Therefore, the second-stage coefficient is positive: defence pacts make for a sharp increase in military cooperation, which makes for a sharp increase in trade.

Table 1.6: The bilateral insecurity reduction

Second stage	(1)	(2)
Estimator:	PPML	
Dependent variable:	Exports	
Bilateral insecurity	-0.346 <sup>a</sup> (0.049)	-0.311 <sup>a</sup> (0.045)
Weak alliance	-0.227 (0.152)	-0.286 <sup>c</sup> (0.149)
RTA	0.040 (0.036)	0.041 (0.033)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.304	167.304
First-stage	OLS	
Estimator:	OLS	
Dependent variable:	Bilateral insecurity	
Defence pact	-1.162 <sup>a</sup> (0.148)	
Common out. def. pacts		-0.076 <sup>a</sup> (0.009)
Weak alliances	-0.213 (0.243)	-0.130 (0.242)
RTA	-0.165 <sup>a</sup> (0.060)	-0.168 <sup>a</sup> (0.060)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.328	167.328

*Note:* OLS, Ordinary Least Square; PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. The panel starts in 1989. Common defence pacts sum all external partners for which country *i* and *j* both have a defence pact. Robust standard errors clustered at country-pair level are in parentheses. Second-stage standard errors are bootstrapped. *a*, *b* and *c* denote significantly different from 0 at the 1%, 5% and 10% level, respectively

## 1.7 General equilibrium and welfare implications

We investigated so far the direct effect of alliances on trade. In this section, we complete our analysis by considering the General Equilibrium (GE) impact – i.e. allowing for changes in aggregated prices, output and expenditures – and derive the welfare gains from alliances. Following the theoretical model presented in section 1.1 and detailed in section 1.B, we can characterise the GE in four equations:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} T_{n \neq s,ij}^{-\theta} F_{n \neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.8)$$

$$\Pi_i^{-\theta} = \sum_j \frac{\tau_{s,ij}^{-\theta} T_{n \neq s,ij}^{-\theta} F_{n \neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\Phi_j^{-\theta}} X_j \quad (1.9)$$

$$\Phi_j^{-\theta} = \sum_i \frac{\tau_{s,ij}^{-\theta} f_{s,ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n \neq s,ij}^{-\theta} F_{n \neq s,ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\Pi_i^{-\theta}} Y_i \quad (1.10)$$

$$\bar{\alpha}_i w_i = \left( \frac{Y_i}{\Pi_i^{-\theta} N_i} \right)^{-\frac{1}{\theta}} \quad (1.11)$$

Equation (1.8) is the structural gravity relation presented in section 1.1 above.  $\Pi_i$  and  $\Phi_j$  are respectively the outward and inward multilateral resistance terms.<sup>61</sup>  $Y_i$  is country  $i$ 's total output and  $\bar{\alpha}_i w_i$  the maximum factory gate price – the marginal cost at the lowest productivity to which a firm can draw in country  $i$ .<sup>62</sup>

In this frame, in keeping with [Arkolakis et al. \(2012\)](#), we can summarise the welfare effect of a change in insecurity costs due to military alliance enforcement in a few parameters. Welfare is defined as the real revenue ( $\frac{Y}{\Phi}$ ). Thus, any change in welfare follows the equality:

$$d \ln(W_j) = d \ln(Y_j) - d \ln(\Phi_j) \quad (1.12)$$

Military alliances contend with bilateral and reciprocal shocks on variable and fixed trade costs sensitive to insecurity. Based on such shocks, we can desegregate country  $j$  changes in welfare as follows:

<sup>61</sup> $\Phi_j$  in eq. 1.10 is obtained by replacing in eq. 1.8  $N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} = \frac{Y_i}{\Pi_i^{-\theta}}$ , which is derived from the market clearance:

$Y_i = \sum_j X_{ij}$ .

<sup>62</sup>Equation 1.11 implicitly assumes that alliances do not affect multilateral insecurity costs ( $\tau_{s,i}$ ). Yet, relaxing this assumption does not impact the GE system. As discussed by [Anderson et al. \(2018\)](#), in an endowment economy the composition of the factory-gate price does not affect the system. In practice, by allowing output-price ratios to adjust to alliances, the GE takes into account any effect that alliances may have on trade or welfare through multilateral insecurity costs. Alternatively, we can directly compute the multilateral insecurity costs as an index of bilateral ones:  $\tau_{s,i} = \sum_j \tau_{s,ij} / I$ . Knowing the trade and alliance elasticities, we can calibrate changes in  $\tau_{s,ij}$ , introduce directly  $\tau_{s,i}$  in the GE and correct the change in real wage. Nonetheless, in section 1.C we further discuss the importance of considering multilateral insecurity and show that in partial equilibrium alliances have a pure bilateral effect.



$$d\ln(W_j) = d\ln(w_j) - \sum_{i=1}^n \psi_{ij} [d\ln(w_i) + d\ln(\tau_{s,ij})] \quad (1.13)$$

Therefore, welfare changes induced by military alliances can be derived from a system of a few parameters: initial trade shares ( $\psi_{ij}$ ), wages ( $w_j$  and  $w_i$ ) and insecurity costs ( $\tau_{s,ij}$ ).<sup>63</sup>

### 1.7.1 GE Empirical application

This General Equilibrium system can be solved using the [Anderson and Yotov \(2016\)](#) methodology and its extension in [Anderson et al. \(2018\)](#). We construct a counterfactual global economy in which no alliance exist - i.e. with the level of international insecurity that would apply in the absence of alliance.<sup>64</sup> By comparing the counterfactual level of exports and real revenues with the baseline, we can assess the GE gains (or losses) in trade and welfare of countries from the observed network of alliances.

We first estimate the partial equilibrium (eq. 1.8)<sup>65</sup> to capture military alliance elasticity and estimates of bilateral trade costs.<sup>66</sup> We develop our analysis for the year 2012 – the last year in our panel. To retrieve baseline multilateral resistance, we estimate the gravity equation for 2012 imposing bilateral trade costs and elasticities from the previous step.<sup>67</sup> We define our counterfactual as the absence of alliance. To obtain the counterfactual multilateral resistance terms, we again estimate the constrained gravity equation setting  $ALL_{ijt} = 0$ . We then determine the endogenous change in output and expenditures:  $\widetilde{X}_i^c = \frac{\widetilde{w}_i^c}{\widetilde{w}_i} X_i$  and  $\widetilde{Y}_i^c = \frac{\widetilde{w}_i^c}{\widetilde{w}_i} Y_i$ . The Change in wage ( $\frac{\widetilde{w}_i^c}{\widetilde{w}_i}$ ) is captured directly by changes in maximum factory gate prices (eq. 1.11).<sup>68</sup> The calibration of trade elasticity  $\theta$  plays a crucial role in the estimation of both prices and welfare. Including tariffs in the standard structural gravity estimation, we directly estimate  $\theta = 3.7$ .<sup>69</sup> Finally, we can quantify the General Equilibrium

<sup>63</sup>Including fixed insecurity costs the welfare equation is:

$d\ln(W_j) = d\ln(w_j) - \sum_{i=1}^n \psi_{ij} [d\ln(w_i) + d\ln(\tau_{s,ij}) + (\frac{1}{\sigma-1} - \frac{1}{\theta})d\ln(f_{s,ij})]$ . Knowing that we have  $d\ln(\tau_{ij}(s)) = \varepsilon_{\ln(\tau_{s,ij})}(ALL_{ijt})$  and  $d\ln(f_{ij}(s)) = \varepsilon_{\ln(f_{s,ij})}(ALL_{ijt})$ , we do not need to identify changes in insecurity costs, but only changes in alliances.

<sup>64</sup>We do not need to specify in this exercise the type of insecurity. Yet, as discussed in section 1.6, this counterfactual insecurity is mainly interpretable as a rise in conflicts with non-state organized actors.

<sup>65</sup>Given that this is a general equilibrium case, intranational trade flows are included. See section 1.3.2 for details on this specification and comparison with our baseline. For consistency and data quality reasons, we use in the GE section TradeProd data only for intra-national flows. We performed a robustness check using TradeProd data for all flows. Results are poorly affected.

<sup>66</sup>Time-varying trade costs are derived from controls, while time-invariant trade costs are captured by country-pair fixed effects. Some exporter-importer fixed effects are dropped due to convergence issues. These missing effects are replaced by regressing for 2012 the estimates of exporter-importer fixed effects on gravity variables and country fixed effects.

<sup>67</sup>Using a PPML estimator, we can directly recover empirical expressions of the multilateral resistance terms ([Fally, 2015](#)). Yet to solve the system, we need to normalise one of the multilateral resistances. We choose to normalise Germany's importer multilateral resistance term so that  $\widetilde{\Phi}_0 = 1$ . Therefore we can derive country  $i$ 's and  $j$ 's multilateral resistance from  $\widetilde{\Pi}_i = (\frac{Y_i X_0}{\exp(\lambda_i)})^{-\frac{1}{\theta}}$  and  $\widetilde{\Phi}_j = (\frac{X_j}{X_0 \exp(\lambda_j)})^{-\frac{1}{\theta}}$ , with  $\widetilde{\lambda}$  the estimated exporter/importer fixed effect.

<sup>68</sup>Given that  $\bar{\alpha}$  and  $N_i$  are fixed, we have  $\frac{\widetilde{w}_i^c}{\widetilde{w}_i} = (\frac{\exp(\lambda_i)^c X_0}{\exp(\lambda_i) X_0^c})^{-\frac{1}{\theta}}$ . Variations in expenditures and outputs trigger new changes in multilateral resistance, which impacts outputs and expenditures and so forth. Translating these variations into changes in exports, we iterate the estimation process until maximum prices converge.

<sup>69</sup>The results of the trade elasticity estimation are reported in table 1.G.13. For robustness reasons, we also perform

effect of military alliances enforced in 2012 as the percentage difference between the baseline and the counterfactual.<sup>70</sup>

## 1.7.2 Results

Results per country are reported in table 1.7.<sup>71</sup> Military alliances improve signatories' exports and welfare, while non-members experience losses. GE effects highly depend on countries' import penetration. The biggest winners in terms of welfare also see sharp increases in their exports, but not the largest ones. These countries are small, developed countries with military alliances with many partners, such as Belgium (16.28%), the Netherlands (+16.20%) and Iceland (+13.12%). We then find a mix of developed and middle-income countries also intensively involved in military alliance treaties, such as Hungary (+11.03%), Mexico (+10.71%), Venezuela (+9.75%), Denmark (+8.75%), Canada (+8.74%) and Argentina (+6.06%). Military alliances are less impactful for large economies, but still bring non-negligible welfare gains: France (+5.34%), Great Britain (+5.21%), Germany (+4.59%) and the USA (+2.42%). On the other hand, the biggest losers are small countries excluded from alliance treaties whose close partners have signed multiple ones. Here, we can cite Ireland (-3.30%), Nigeria (-2.85%), Malta (-2.44%), Cyprus (-2.35%) and Sweden (-2.21%). These negative effects are not the result of a rise in insecurity for non-signatory countries, which is assumed as constant,<sup>72</sup> but simply the consequence of a general equilibrium trade diversion mechanism.<sup>73</sup> In addition, our results are robust to the inclusion of heterogeneous alliances and trade elasticities.<sup>74</sup>

## 1.7.3 Scenarios analysis

International tensions have been exacerbated in recent years. The war in Ukraine has brought conflict back to Europe with new threats and strategic interests that are shaking the post-Cold War balance.

the general equilibrium analysis using two alternative calibrations of  $\theta$ . In keeping with [Head et al. \(2014\)](#) and [Melitz and Redding \(2013\)](#) we calibrated  $\theta = 4.25$ , and in keeping with [Anderson et al. \(2018\)](#)  $\theta = 6$ . Mechanically, the higher  $\theta$ , the smaller the welfare variations. Nonetheless, our conclusions are robust to these alternative calibrations of  $\theta$ .

<sup>70</sup>Changes in exports and welfare are simply:  $\frac{X_j}{X_j^c} = \frac{\sum_j X_{ij} - \sum_j \bar{X}_{ij}^c}{\sum_j \bar{X}_{ij}^c}$  and  $\frac{W_j}{W_j^c} = \frac{Y_j / \bar{\Phi}_j - \bar{Y}_j^c / \bar{\Phi}_j^c}{\bar{Y}_j^c / \bar{\Phi}_j^c}$ .

<sup>71</sup>Minor additional sensitivity checks are performed. We replicate the analysis while controlling for globalisation as in table 1.G.2 column 8. We also differentiate between defence pacts and weak alliances, removing only defence pacts. In both checks, results are barely affected.

<sup>72</sup>We properly discuss this assumption in section 1.4.

<sup>73</sup>Note that losers may experience (very small) gains in exports. This is the result of an increase in their inward and outward multilateral resistance terms associated with a drop in their factory gate prices. The increase in expenditure by winners causes non-signatory countries to see trade diverted from their internal market to the winners. This diversion offsets (or overcomes) the decrease in their exports due to the fall in output.

<sup>74</sup>We showed in section 1.3.3 that, given that trade elasticities are not constant, alliances have a country-pair specific effect. We adjust our GE analysis to include bilateral trade elasticities. We allow trade elasticities to be country-pair specific and the price-index elasticity to be country specific. Theoretically, the above-presented GE structure and welfare system still hold. After estimating the average effect on military alliances, we introduce the distortion in coefficients derived from the quantile estimation in section 1.3.3. Then, we determine the price-index elasticities. We calibrate the average elasticity at 3.7 and apply the country-specific distortion. In keeping with our model extension in section 1.B.3, a good proxy can be directly derived from the importer's average trade elasticities. We present the new GE results in appendix table 1.G.14. On the whole, the introduction of bilateral elasticities tempers the welfare variations. Nonetheless, our conclusions are barely affected.

The future may bring a substantial reshaping of military alliances, especially among European countries. Using our general equilibrium approach, this section analyses the ramifications on countries' welfare of potential major alliance network disruption.<sup>75</sup> We define three scenarios: i) the expansion of NATO to neutral European countries and partner nations, ii) NATO breaking with Eastern countries, and iii) the creation of a new Eastern block.<sup>76</sup> Extending the network of alliances, the first scenario induces a pure reduction in international insecurity. On the contrary, the second scenario captures a pure increase in insecurity, while the third scenario assumes a mix of bilateral insecurity changes.<sup>77</sup>

Results per country are presented in the appendix table 1.G.15. Despite the assumed important change in the alliance network, any scenario's GE effect is insufficient to induce a major impact on large or unconcerned economies. For countries like the US, France, UK and Germany, maintaining NATO in the East or expanding the treaty to their closest partners has noticeable, but not major economic outcomes. Furthermore the changes in the GE trade diversion are weak and poorly affect neutral economies. On the other hand, for countries targeted as potential switchers, the choice to switch would have drastic repercussions. Leaving NATO would severely reduce their real revenue, which a new Eastern block could temper but not offset. Moreover, for still-neutral countries – especially in Europe – joining NATO (i.e. the closest and largest alliance network) would bring remarkable welfare gains. For example, Bulgaria would increase its real revenue by 10.27%, Finland by 11.51%, Sweden by 15.73%, Switzerland by 17.13% and Ireland by 24.58%.

## 1.8 Conclusion

This chapter provides a systematic analysis of the impact of military alliances on trade using a panel of 6,972 country pairs from 1967 to 2012. Taking a structural gravity approach, we show that military alliances have a strong positive causal effect on bilateral exports. Namely, our baseline specification shows that the enforcement of a military alliance engenders an increase of 60%. We perform numerous sensitivity tests and show that results are robust to a variety of alternative specifications. However, this average effect is highly heterogeneous across country pairs, depending on the nature of the treaties and the economic size of partners. Furthermore, an instrumental variable approach and a DATT approach confirm the causal interpretation of our results. We confirm the validity and prevalence of the insecurity

<sup>75</sup>We do not observe Russia or Ukraine's intranational trade, which excludes them from the analysis. Therefore, we do not aim to study the impact of war on the two belligerents directly, but the ramifications of new military relationships between countries indirectly involved.

<sup>76</sup>We assume that i) Sweden, Finland, Japan, Austria, Australia, Ireland, South Korea, New Zealand, Colombia, Malta, Switzerland, Pakistan and Serbia-Montenegro become full NATO members, ii) Albania, Bulgaria, China, Finland, Sweden, Hungary, Pakistan, Poland, Romania, Serbia-Montenegro and Turkey leave NATO or terminate any military alliance with its members, and iii) that after leaving NATO, they form one common military alliance together.

<sup>77</sup>We argued in the present chapter that overall alliances only affect signatory country-pairs. Yet, we also showed that in specific contexts, such as the Cold-War, some change in trade with non-signatory partners may be observed (see appendix section 1.E). Allowing for such indirect changes in the scenarios tempers the positive effects of alliances but does not affect our conclusions.

mechanism: military alliances increase trade by reducing bilateral insecurity. Finally, we perform a general equilibrium analysis to quantify the welfare effect of alliances. Building a counterfactual for 2012, we show that intensive involvement in the signature of military alliances brings significant welfare gains while being neutral induces marked losses. We then analyse different scenarios to demonstrate that reshaping the military alliance network in response to the war in Ukraine could have considerable welfare ramifications on the concerned economies.

Our findings have important scientific and policy implications. First, they point up the need to consider the specific role of security and international military relations to understand trade and globalisation. Second, they show the efficiency of military alliances, particularly defence pacts, at guaranteeing the safety and inter-state cooperation required for economic agents to trade. That opens an interesting research agenda. The important macro-economic effect of alliances calls for a better understanding of the micro-level dynamic. This questions the heterogeneity in the exposure to insecurity among sectors and firms, as well as the capacity of international security cooperation policies to address its complexity. Nonetheless, the unambiguous welfare gains that alliances bring their members should give policymakers strong incentives to promote the signature of such treaties. Although some may fear that they create relations of domination between nations, our findings suggest on the contrary that they bring more favourable welfare gains to small economies.

Table 1.7: GE exports and welfare gains from alliances (2012)

Country (Iso3)	Exports	Real revenues	Country (Iso3)	Exports	Real revenues
ALB	2.34	-0.38	ISL	29.96	13.12
ARG	21.54	6.06	ISR	-2.66	-1.84
AUS	4.16	0.46	ITA	35.10	3.53
AUT	-1.62	-1.93	JPN	7.37	0.40
BEL	19.95	16.28	KEN	-5.54	-0.34
BGD	-0.36	-0.99	KOR	5.45	0.61
BGR	9.02	1.68	LKA	-2.06	-0.91
BOL	29.58	11.60	MAR	-0.15	-0.52
BRA	20.02	0.74	MEX	28.92	10.71
CAN	44.41	8.74	MLT	1.02	-2.44
CHE	-1.05	-1.89	MYS	0.63	-0.71
CHL	11.76	6.11	NGA	1.44	-2.85
CHN	-1.14	-0.13	NLD	17.80	16.20
CIV	13.14	0.30	NOR	20.31	6.61
CMR	0.20	-1.25	NZL	-0.51	-0.44
COL	38.52	3.32	PAK	25.09	5.19
CYP	-1.04	-2.35	PER	25.07	2.79
DEU	26.66	4.59	PHL	4.21	1.70
DNK	26.85	8.75	POL	61.18	5.11
DZA	-0.45	-1.02	PRT	56.90	6.94
ECU	31.62	6.26	ROM	2.24	-0.19
EGY	-3.64	-0.68	SAU	0.38	-0.96
ESP	45.81	4.14	SER	-5.52	-0.55
FIN	-2.55	-1.28	SGP	-0.39	-0.38
FRA	45.15	5.34	SWE	-2.32	-2.21
GAB	1.38	3.66	THA	0.31	-0.64
GBR	41.19	5.21	TUN	0.22	-0.77
GRC	51.44	4.98	TUR	60.8	2.53
HKG	0.02	-1.36	URY	37.82	3.86
HUN	26.82	11.03	USA	65.90	2.42
IDN	0.14	-0.32	VEN	18.17	9.75
IND	-2.36	-0.20	VNM	0.05	-0.45
IRL	1.25	-3.30			

*Note:* The real revenue is our measure of welfare. All numbers are percentage variations.

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# Appendix

## 1.A Military alliance treaties

### 1.A.1 Overview

Military alliances are international treaties designed to develop international security cooperation policies. As defined by Gibler (2008) military alliances can be divided into four categories depending on the degree of restriction and involvement of signatories. First, the *military entente* implies a diplomatic exchange of information among members before taking any military decision. Second, the *neutrality pact* specifies that signatories must stay neutral in the event of a conflict involving one party to the pact. Third, the *non-aggression pact* states that signatories cannot declare war or engage in military action against treaty members. Fourth, the *defence pact* is where signatories agree on collective but centralised military management. It does not deny members' sovereignty, but enforces strong security cooperation in areas that matter to the signatory countries. Therefore, defence pacts reach a highly specific level of cooperation. While the first three categories of alliance mainly describe different international policies to keep peace between members, the defence pacts imply military cooperation to protect signatories from any common threats.

### 1.A.2 Defence pacts examples

The most famous defence pact is the *North Atlantic Treaty* (NATO). Signed on the fourth of April 1949 and still in force, it concerns most North American and Western European countries and was later expanded to a number of Eastern European nations. Created to ensure protection against the USSR and its satellites, it implies strong and centralised military cooperation. It was also designed with important economic and institutional objectives in mind such as economic collaboration, free institutions and stable well-being.<sup>78</sup> Yet, to pursue these objectives, the treaty specifies only military

<sup>78</sup>"Article 2. The Parties will contribute toward the further development of peaceful and friendly international relations by strengthening their free institutions, bringing about a better understanding of the principles upon which these institutions are founded, and promoting conditions of stability and well-being. They will seek to eliminate conflict in their international economic policies and will encourage economic collaboration between any or all of them."

cooperation policies. Moreover, the ninth article includes the creation of a central council in charge of compliance with the alliance's constraints, organisation and objectives.

The *Treaty on Collective Security* (TCS) was signed by former USSR states in 1992 and is still in force. It was signed in recognition of the inability of the Commonwealth of Independent States to provide the required economic and commercial prosperity among members despite the tariff liberalisation it includes. The TCS aims to achieve members' trade objectives by enabling lasting stability and security throughout the region due to common and centralised management of military matters. It also specifies the creation of a collective security council in charge of defence decisions, armed forces coordination and the application of the treaty's purposes.

The *Joint Defense and Economic Cooperation Treaty* between the States of the Arab League was created in 1951 by Egypt, Lebanon, Saudi Arabia, Syria, Yemen Arab Republic and Iraq, and joined later by a number of Arab states. Even though the treaty has encountered problems due to internal tensions over members' relations with Israel, it is still in force and was reinvigorated following the USA intervention in Iraq. Once again, it concerns close military cooperation and the creation of a centralised council. It also contains explicit economic and trade objectives and aims to favour the development and trade of signatory countries.<sup>79</sup> Moreover, to assist the first council, a second council in charge of economic issues was created. Yet, this council cannot propose or enforce standard trade liberalisation policies such as tariff reductions.

The *Defense Pact of the African and Malagasy Union* was signed in 1961 by twelve French-speaking sub-Saharan countries to protect themselves from both internal and external interference. Nonetheless, the pact did not last long since it was terminated in 1964. The core of the treaty was the enforcement of peace and stability in the region and military cooperation by members. Yet, it also introduced a mandatory contribution to the development of free institutions, well-being and economic collaboration. Similarly to the previous example, international security was also considered here as a necessity for economic development. Moreover, the treaty stipulated the creation of a central council to take decisions with a two-thirds majority of pact members regarding the alliance's procedures.

The last example of defence pact is the *Treaty Instituting the Arab-Maghreb Union*. It was signed in 1989 by North African and Arab countries and is still in effect today. The agreement covers common management of defence and stability matters, the creation of a presidential council to centralise decisions and a judicial body to ensure the legal enforcement of decisions. A striking point in our case is the explicit objective of trade liberalisation and free movement of persons, services, capital and

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<sup>79</sup>"Article 7. In order to fulfil the aims of this Treaty and to bring about security and prosperity in Arab countries and in an effort to raise the standard of life therein, the contracting States undertake to collaborate for the development of their economic conditions, the exploitation of their natural resources, the exchange of their respective agricultural and industrial products, and generally to organise and coordinate their economic activities and to conclude the necessary inter-Arab agreements to realise such aims."

goods.<sup>80</sup> The treaty's third article also includes a clear objective of common economic and industrial development for members.

### 1.A.3 Weaker alliances examples

A first example is the neutrality pact between Chile and Argentina called the *Treaty of Peace and Friendship between Chile and Argentina*. It was driven by increasing border disputes between the two states since 1970 and negotiations to settle them. It was finally signed in 1984 and is still in force today. It lays down military and economic objectives, especially maritime goals, and the establishment of a commission. Yet, contrary to a defence pact, it does not imply close cooperation between states but merely a frame to avoid armed conflicts and agree on each state's sovereignty.

The *Treaty of Peace, Friendship and Cooperation between India and the Soviet Union*, coded as a non-aggression pact, was signed in 1971 and was in effect until 1991. The treaty provides for respect for members' sovereignty and borders and an absence of interference in any domestic affairs. Moreover, it expresses the importance of economic cooperation and trade. Yet, it provides for no supra-national institution to be created. The agreement provides for a guarantee from participants, but neither centralised military cooperation nor the ambitions of a defence pact.

The *treaty of Friendship and Co-operation between France and Russia*, an entente pact, was signed by Francois Mitterrand and Boris Yeltsin in 1992. It is still in force today without any objection having ever been raised by either state. The main goal of the agreement is to ensure immediate consultation between signatories in the event of security issues or important diplomatic decisions. It also stipulates institutional and economic objectives such as the development of the manufacturing sectors, the promotion of democratic institutions and the facilitation of the movement of capital, persons and goods. Yet, there is no further military cooperation, no guarantee of peace, and no supra-national institution to ensure the application of the treaty.

## 1.B Theoretical model

### 1.B.1 Standard model

We start with a constant elasticity of substitution utility function. Consumers in each country maximise their utility by consuming  $q(\omega)$  units of each differentiated good  $\omega$ , noted by the following function:

$$U = \left[ \int_{\Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right]^{\frac{\sigma}{\sigma-1}} \quad (1.14)$$

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<sup>80</sup>"Article Two. The Union aims at:(...) - Working gradually towards achieving free movement of persons and transfer of services, goods and capital among them."

where  $q$  is the consumed quantity, and  $\sigma$  the elasticity of substitution between two varieties of goods. After maximisation under the revenue constraint, we can define the consumed quantity for a specific variety as:

$$q(\omega) = p(\omega)^{-\sigma} X \left( \int_{\Omega} p_{\omega''}^{1-\sigma} \right)^{\frac{\sigma-1}{1-\sigma}} \quad (1.15)$$

and firm  $\omega$ 's revenue:

$$x(\omega) = X_j \left( \frac{p(\omega)}{P_j^{1-\sigma}} \right)^{1-\sigma} \quad (1.16)$$

where  $x(\omega) = q(\omega)p(\omega)$ ,  $X_j$  the total revenue of country  $j$ , and  $P_j = \int_{\Omega} p_{\omega''}^{1-\sigma}$  the Dixit-Stiglitz price index. Thus, the total consumed value in country  $j$  of goods variety  $\omega$  can be understood as the share of the country's total revenue allocated to the consumption of variety  $\omega$ .

Next, we consider firm productivity level  $\gamma$  such that marginal cost  $\alpha = 1/\gamma$ . In keeping with [Helpman et al. \(2008\)](#), we assume that the distribution of firm productivity  $\gamma$  in each country follows a Pareto distribution  $G$  with:

$$G_i(\alpha) = (\alpha^\theta - \underline{\alpha}^\theta) / (\bar{\alpha}_i^\theta - \underline{\alpha}^\theta) = \frac{\alpha^\theta}{\bar{\alpha}_i^\theta} \quad (1.17)$$

where  $\theta$  is the parameter determining the shape of the distribution,  $\alpha = \frac{1}{\gamma}$  the firm's marginal cost,  $\bar{\alpha}_i$  the maximum marginal cost (or minimum productivity) to produce in country  $i$ , and  $\underline{\alpha}$  the minimum marginal cost (or maximum productivity). Therefore, naming  $N$  the number of firms and considering solely exports from country  $i$  to country  $j$  we have:

$$x_{ij}(\alpha) = X_j \left( \frac{p(\alpha)}{(\sum_l N_l \int_0^{\bar{\alpha}_l} p_{lj}(\alpha)^{1-\sigma} dG(\alpha))^{\frac{1}{1-\sigma}}} \right)^{1-\sigma} \quad (1.18)$$

Firms present monopolistic competition and a CES demand function. Therefore, by introducing insecurity costs – discussed in section 1.1 – and other sorts of trade costs, we obtain the following price and profit functions for each variety:

$$p_{ij}(\alpha) = \frac{\sigma}{\sigma-1} w_i T_{n,ij} \alpha \quad (1.19)$$

$$\pi_{ij} = \left( \frac{x_{ij}(\alpha)}{\sigma} \right) - F_{n,ij} \quad (1.20)$$

where  $T_{n,ij} = \Pi^n \tau_{n,ij}$  is a product of variable trade costs with  $n$  the  $n$  potential source of iceberg costs, including  $\tau_{s,ij}$  the insecurity cost, but also all variable trade costs sensitive to other parameters (geography, standard trade policies, institutions, etc.), and  $F_{n \neq s,ij}$  a vector of fixed trade costs

insensitive to insecurity that firms have to pay to enter country  $j$  from  $i$ .

Once all exports from  $i$  to  $j$  have been aggregated, taking into account equations (1.22) and (1.23), and setting  $\underline{\alpha} = 0$  for solving issues<sup>81</sup>, we can define total exports from  $i$  to  $j$ :

$$X_{ij} = X_j \frac{N_i w_i^{1-\sigma} V_{ij} T_{n,ij}^{1-\sigma}}{\sum_l N_l w_l^{1-\sigma} V_{lj} T_{n,lj}^{1-\sigma}} \quad (1.21)$$

where  $V$  is defined as in Helpman et al. (2008):

$$V_{ij} = \int_0^{\alpha_{ij}^*} \alpha^{1-\sigma} dG(\alpha) \quad (1.22)$$

$\alpha_{ij}^*$  is by definition the level of productivity for which the profit from exporting,  $\pi_{ij}$ , is zero:

$$\pi_{ij} = \left( \frac{x_{ij}(\alpha^*)}{\sigma} \right) - F_{n \neq s, ij} = 0 \quad (1.23)$$

where:

$$\alpha^* = (\sigma - 1) \sigma^{\frac{\sigma}{\sigma-1}} \left( \frac{X_j}{P'_l F_{n \neq s, ij}} \right)^{\frac{1}{\sigma-1}} \frac{1}{w_i T_{n,ij}} \quad (1.24)$$

with index price

$$P'_l = \sum_l N_l \int_0^{\alpha_{lj}^*} p_{lj}(\alpha)^{1-\sigma} dG(\alpha) \quad (1.25)$$

Once equation (1.27) is plugged into equation (1.25), we can develop V. Combined with equation (1.24), we obtain the following expression for bilateral exports:

$$X_{ij} = X_j \frac{N_i (\bar{\alpha}_i w_i)^{-\theta} T_{n,ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]}}{\sum_l N_l (\bar{\alpha}_l w_l)^{-\theta} T_{n,lj}^{-\theta} F_{n \neq s, lj}^{-[\frac{\theta}{\sigma-1}-1]}} \quad (1.26)$$

Finally, after defining the importer multilateral resistance term:

$$\Phi_j = \left( \sum_l N_l \bar{\alpha}_l^{-\theta} w_l^{-\theta} T_{n,lj}^{-\theta} F_{n \neq s, lj}^{-[\frac{\theta}{\sigma-1}-1]} \right)^{-\frac{1}{\theta}} \quad (1.27)$$

we obtain the structural gravity equation:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} T_{n,ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.28)$$

Or outlying insecurity costs from trade cost aggregates:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s,ij}^{-\theta} T_{n \neq s, ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.29)$$

<sup>81</sup>This implies that we assume that there is always a firm that is productive enough to export at least a small amount.



with  $T_{n \neq s, ij}$  is the variable trade costs sets excluding insecurity costs  $\tau_{s, ij}$ .

### 1.B.2 Extension with fixed insecurity costs.

In the standard model, we assume that insecurity costs are equivalent to an expropriation risk which can be model as a variable trade cost. Yet, insecurity costs may also be related to barriers inducing a fixed cost (controls, procedures and information requirements).

We introduce insecurity barriers in our model. Firms need to address expensive procedures and information requirements to lift these barriers and enter the foreign market. The higher the bilateral insecurity, the higher the barriers and therefore the higher the cost to lift them. This is directly interpretable as a fixed trade cost sensitive to insecurity: to enter the market, firms pay a cost that depends on the degree of bilateral insecurity but does not vary with the exported quantity. So it does not affect the price function, but directly affects profit:

$$\pi_{ij} = \left( \frac{x_{ij}(\alpha)}{\sigma} \right) - F_{n, ij} \quad (1.30)$$

$x_{ij}(\alpha)$  is the firm's revenue function and  $F_{n, ij}$  a vector of fixed trade costs that firms have to pay to enter country  $j$  from  $i$ , including  $f_{s, ij}$ , the fixed insecurity cost derived from insecurity barriers, but also all fixed trade costs sensitive to other sources  $n$ . If the market entry cost decreases, the firm's profit increases. When firms switch from negative profit to positive profit, they start exporting, which increases the number of varieties sold from  $i$  to  $j$ .

Therefore, we can derive at aggregated level the following structural gravity equation outlying the insecurity trade costs  $\tau_{s, ij}$  and  $f_{s, ij}$ :

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \frac{X_j}{\Phi_j^{-\theta}} \tau_{s, ij}^{-\theta} f_{s, ij}^{-[\frac{\theta}{\sigma-1}-1]} T_{n \neq s, ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.31)$$

$X_{ij}$  is the total exports from country  $i$  to country  $j$ ,  $N_i$ , the number of firms in the exporting country,  $\bar{\alpha}_i$ , the maximum marginal cost of country  $i$ 's technology, and  $w_i$  the wage in country  $i$ 's economy.  $X_j$  is the total revenue of country  $j$ , and  $\Phi_j$ , the importer's multilateral resistance term, while  $T_{n \neq s, ij}$  and  $F_{n \neq s, ij}$  are respectively the variable and fixed trade costs sets, but excluding the insecurity costs. Again, like other trade costs, insecurity costs ( $\tau_{s, ij}$  and  $f_{s, ij}$ ) have negative elasticities.<sup>82</sup> Hence, by reducing both variable and fixed insecurity costs, the enforcement of a military alliance between countries  $i$  and  $j$  increases bilateral exports  $X_{ij}$ .<sup>83</sup>

<sup>82</sup>These elasticities depend both on  $\theta$ , the Pareto shape parameter of the firms' productivity distribution. Yet, fixed trade cost elasticity also depends on  $\sigma$ , the elasticity of substitution. Here, we assume  $\theta > \sigma - 1$ . Otherwise, fixed trade costs elasticity is positive.

<sup>83</sup>We do not make any composition assumptions. Military cooperation policies described in appendix 1.A suggest that both variable and fixed insecurity costs are reduced by alliances. In our analysis of the overall effect of alliances and

### 1.B.3 Extension with bilateral trade elasticities

A way to deal theoretically with the constant trade elasticity issue is to rule out the CES assumption in favour of other utility functions consistent with the idea of sub-convex gravity (Mrázová and Neary, 2017). Yet, the CES utility function has the advantage of being simple to use, works well with Pareto productivity distributions, and allows for a structural gravity equation with heterogeneous firms (Chaney, 2008). We propose another solution to capture bilateral trade elasticities: relaxing the assumption of constant technology in our standard model.

We start with the same structure as earlier, that is to say, the CES demand function, monopolistic competition and firm heterogeneity with a Pareto distribution of productivity. Yet, we allow the Pareto shape parameter  $\theta$  to be exporter specific and assume that worldwide technology has been developed to match the sub-variety of dominant markets. Precisely, we apply a distortion  $\zeta_j$  on the productivity distribution. In keeping with empirical evidences presented in Carrère et al. (2020) and section 1.3.3, we assume  $\frac{\delta\theta_i}{\delta\frac{Y_i}{\pi_i^{\theta_i}}} < 0$  and  $\frac{\delta\zeta_j}{\delta\frac{X_j}{\phi_j^{\theta_j}}} < 0$ . The larger countries  $i$ 's and  $j$ 's share of global revenue, the greater the firm's probability of drawing high productivity. In other words, firms still have a monopoly on their variety, but produce a sub-variety for each destination that has its own productivity level and a greater chance of being more productive in sub-varieties designed for the largest markets. For example, a firm might produce blue T-shirts (its variety) with the highest level of productivity for the sub-variety designed for the US (chemical dyed large blue T-shirts). However, it would have to divert from this optimal production chain when producing the sub-variety for Sweden (naturally dyed long blue T-shirts), where consumer characteristics and preferences are different.

We define the country-pair specific productivity distribution as follows:

$$G_{ij}(\alpha) = (\alpha^{(\theta_i\zeta_j)} - \underline{\alpha}^{(\theta_i\zeta_j)}) / \overline{\alpha}'^{(\theta_i\zeta_j)} - \underline{\alpha}^{(\theta_i\zeta_j)} = \frac{\alpha^{(\theta_i\zeta_j)}}{\overline{\alpha}'^{(\theta_i\zeta_j)}} \quad (1.32)$$

with  $\overline{\alpha}' = \frac{X_j}{\phi_j} \overline{\alpha}$  the maximum marginal cost. Therefore, the firms' probability of drawing sub-variety productivity  $\gamma_{ij}$  and the productivity distribution in the economy are determined by  $\theta_i$  and  $\zeta_j$ . Each country-pair presents a specific productivity distribution, while each firm draws different outputs depending on the destination market. Nonetheless, because each country-pair's productivity is Pareto distributed, the observed average productivity of firms in a given origin country is also Pareto distributed.<sup>84</sup> In this extended model, the price and profit equations are unchanged. Yet, because the

---

the role of bilateral insecurity, we recognise for interpretation reasons the existence of both variables in fixed insecurity costs, but do not need to disentangle them.

<sup>84</sup>At country-pair level, many firms draw small productivity while a few firms draw high productivity. Taking the firms' average productivity, we observe the same thing. Many firms draw small productivity for all destinations while a few firms draw only high productivity. Nonetheless, in this framework, some firms that are on average not very productive may have a low level of productivity for the large majority of destinations, but a high level of productivity for a few destinations.

productivity distribution is different from the standard model, the firms' productivity is affected and their prices too.

Running through the model, we obtain the following expression of the structural gravity equation with bilateral trade cost elasticities:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta_i \zeta_j} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_i \zeta_j} w_i^{-\theta_i \zeta_j} \frac{X_j}{\Phi_j'} \tau_{s,ij}^{-\theta_i \zeta_j} f_{s,ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1}-1]} T_{n \neq s, ij}^{-\theta_i \zeta_j} F_{n \neq s, ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1}-1]} \quad (1.33)$$

with

$$\Phi_j' = \sum_l N_l \bar{\alpha}_l^{-\theta_l \zeta_j} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_l \zeta_j} w_l^{-\theta_l \zeta_j} T_{n, lj}^{-\theta_l \zeta_j} F_{n, lj}^{-[\frac{\theta_l \zeta_j}{\sigma-1}-1]} \quad (1.34)$$

This distorted structural gravity equation raises new trade cost elasticities: variable trade cost elasticity  $-\theta_i \zeta_j$  and fixed trade cost elasticity  $-\left[\frac{\theta_i \zeta_j}{\sigma-1} - 1\right]$ . The lower  $\theta_i$  and  $\zeta_j$ , the lower the trade elasticities. Therefore, large economies are less affected by insecurity costs. This explains why we observe alliances elasticities inversely proportional to the value of trade (cf. section 1.3.3).

## 1.C Multilateral insecurity

In section 1.6, we showed that alliances strongly increase trade by reducing bilateral insecurity. Yet, in theory, we may expect alliances to also induce multilateral changes in insecurity. In keeping with (Martin et al., 2008), we can define multilateral insecurity as a exporter-specific marginal cost  $\tau_{s,i}$  leading to the price function:

$$p_{ij}(\alpha) = \frac{\sigma}{\sigma-1} w_i \tau_{s,i} \tau_{s,ij} T_{n \neq s, ij} \alpha \quad (1.35)$$

and, deriving our model,<sup>85</sup> to the gravity equation:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta} w_i^{-\theta} \tau_{s,i}^{-\theta} \frac{X_j}{\Phi_j'} \tau_{s,ij}^{-\theta} T_{n \neq s, ij}^{-\theta} F_{n \neq s, ij}^{-[\frac{\theta}{\sigma-1}-1]} \quad (1.36)$$

In this framework, alliances could increase bilateral trade  $X_{ij}$  by reducing the bilateral insecurity cost  $\tau_{s,ij}$  and the multilateral one  $\tau_{s,i}$ . In our baseline estimation, any change in  $\tau_{s,i}$  are captured by the exporter-time fixed effects. Alliances are therefore restricted to affect trade only through bilateral insecurity. Then, in the General Equilibrium approach, we implement a bilateral insecurity shock resulting from alliances, and, under several assumptions,<sup>86</sup> allow multilateral insecurity to adjust. These methods consider the existence of multilateral insecurity, but do not permit to observe whether they are affected by alliances.

<sup>85</sup>see the appendix section 1.B for details)

<sup>86</sup>Mainly, we do not need to disentangle multilateral insecurity from other marginal costs, but must assume that multilateral insecurity changes are either redistributed in revenues or compensated by a proportional change in wage.

To address this issue, we estimate below the alliances' effect on trade allowing for changes in both bilateral and multilateral insecurity. In order to allow for a change in multilateral insecurity, we cannot include exporter-time fixed effects.<sup>87</sup> However, this may cause omitted variable bias, since changes in the exporter output or price-index may influence the exporter's willingness to sign alliances. Therefore, we use the instrumental variable detailed in section 1.4. We present the results in table 1.C.1. In column one, we use a two-stage OLS-PPML strategy and in column two a control function (Lin and Wooldridge, 2019). First-stage coefficients are not affected by the absence of exporter-time fixed effect.<sup>88</sup> Furthermore, using the *plausible exogeneity test* proposed by Conley et al. (2012), we confirm that the IV still sufficiently respect the exclusion restriction (see table 1.C.2). Finally, in the second stage, alliances' coefficient is very weakly affected, while capturing a reduction in multilateral insecurity should have increased the estimated elasticity. This supports that alliances' impact is purely bilateral.

Table 1.C.1: Allowing for bilateral and multilateral insecurity changes

Dependent variable:	exports	exports
Second stage	IV	Ctrl. function
Estimator:	PPML	PPML
Alliance	0.647 <sup>a</sup> (0.024)	0.614 <sup>a</sup> (0.026)
RTA	0.198 <sup>a</sup> (0.017)	0.193 <sup>a</sup> (0.017)
First stage residuals		-0.198 <sup>a</sup> (0.057)
Exporter x Year FE	no	no
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666
First-stage		
Estimator:	OLS	OLS
Dependent variable:	Alliance	Alliance
Common out. alliances	0.056 <sup>a</sup> (0.007)	0.056 <sup>a</sup> (0.007)
RTA	0.013 <sup>c</sup> (0.007)	0.013 <sup>c</sup> (0.007)
Exporter x Year FE	no	no
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.712	320,712
KPW F-stat	68	68
KPW LM-stat	11	11

*Note:* OLS, Ordinary Least Squares; PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. Standard errors clustered at the exporter and importer levels are in parentheses. Second-stage standard errors are bootstrapped. a, b and c denote significantly different from 0 at the 1%, 5% and 10% level, respectively.

<sup>87</sup>Changes in bilateral insecurity can also affect the exporter price index. Removing the exporter-time fixed effects we also allow alliances to affect trade through this channel.

<sup>88</sup>See table 1.4 for comparison.

Table 1.C.2: Plausible exogeneity test, multilateral insecurity

Dep var: exports	<i>Union of Confidence Interval estimations</i>		
Instrumented var.	$\nu$	Min	Max
	-0.030	95% CI	95% CI
Alliance		0.448	1.166

*Note:* UCI based on the IV's  $\nu$  coefficient from a regression of exports on interest variables and the IV

## 1.D Two-way (robust) fixed effects estimations

As pointed up by [De Chaisemartin and d'Haultfoeuille \(2020\)](#), heterogeneous treatments or treatment effects over time and groups may return false results in the case of two-way fixed effects estimations. Comparing groups (here country-pairs) that are not treated at the same time or that experience different outcomes following the treatment could cause negative weights in the (bias) ATE. The estimator developed by [De Chaisemartin and d'Haultfoeuille \(2020\)](#) is sufficient to deal with this bias. But, because we want to address the effect's dynamic, we use the later estimator developed by [De Chaisemartin and d'Haultfoeuille \(2024\)](#), which provides event study results.

The intuition behind this estimation is that avoiding these negative weights entails a comparison of the first-time switchers' t-1 to t+l outcome evolution with the t-1 to t+l outcome evolution of country-pairs whose treatment has hitherto remained stable; with t the treatment time and l the event time. Our panel is balanced, so we estimate the treatment effect from positive switchers (i.e. alliance signatures):

$$\delta_l = \frac{DID_{+,l}}{DID_{+,l}^D} \quad (1.37)$$

where:

$$DID_{+,l} = \sum_{g:D_{g,1}=0, t_g < T_u - l} \frac{N_{g,t_g+l} \beta^{t_g+l}}{N_l^1} \left[ (Y_{g,t_g+l} - Y_{g,t_g-1}) - \sum_{g:D_{g',1}=0, t_{g'} > t_g+l} \frac{N_{g',t+l}}{N_{t_g+l}^u} (Y_{g',t_g+l} - Y_{g',t_g-1}) \right] \quad (1.38)$$

and

$$DID_{+,l}^D = \sum_{g:D_{g,1}=0, t_g < T_u - l} \frac{N_{g,t_g+l} \beta^{t_g+l}}{N_l^1} \left[ (D_{g,t_g+l} - D_{g,t_g-1}) - \sum_{g:D_{g',1}=0, t_{g'} > t_g+l} \frac{N_{g',t+l}}{N_{t_g+l}^u} (D_{g',t_g+l} - D_{g',t_g-1}) \right] \quad (1.39)$$

with  $t_g$  the time of group  $g$ 's treatment variation,  $\beta \in (0, 1]$  the planner's discount rate,  $D$  the treatment variation,  $T_u$  the last observed period with a group untreated since period 1,  $N_l^1 = \sum_{g:D_{g,1}=0, t_g < T_u - l} N_{g,t_g+l} \beta^{t_g+l}$  the discounted number of units in groups reaching  $l$  periods after their first treatment or before  $T_u$ , and  $Y$  the outcome – in our case exports. Yet, since  $ALL_{ijt}$  is a dummy

variable,  $DID_{+,l}^D = 1$ . Therefore, we have:

$$\delta_l = DID_{+,l} \quad (1.40)$$

We present the results for military alliances in table 1.D.1 and defence pacts in table 1.D.2. They show a growing dynamic effect and confirm the robustness of our identification strategy.

However, De Chaisemartin and d’Haultfoeuille (2024) estimator is linear which in a gravity setting may bias results (Silva and Tenreyro, 2006). Therefore, we perform an alternative two-way (robust) fixed effects approach developed by Nagengast and Yotov (2023); Wooldridge (2023) which allows for ppml estimator. The intuition is similar. By interacting each cohort with the post-treatment periods, it directly addresses the negative weights issue that may results from heterogeneous effects. The results, displayed on graph 1.D.1, are similar to the linear estimation and support our conclusions.

Table 1.D.1: Military alliances: Chaisemartin D’Haultfoeuille (2020) estimator

Dependent variable: Bilateral exports						
Time	t	t+1	t+2	t+3	t+4	t+5
Alliance’s coeff	0.339 <sup>a</sup>	0.457 <sup>a</sup>	0.500 <sup>a</sup>	0.581 <sup>a</sup>	0.661 <sup>a</sup>	0.682 <sup>a</sup>
Standard error	(0.085)	(0.086)	(0.090)	(0.093)	(0.093)	(0.094)
RTA control	yes	yes	yes	yes	yes	yes
No. observ.	128.004	127.878	127.788	127.594	121.022	114.458
No. switchers	250	248	248	246	242	238

*Note:* Dependent variable is the inverse hyperbolic sine transformation of exports from country i to country j at time t in millions of current dollars. t is the year of the pair’s alliance signature. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.D.2: Defence pacts: Chaisemartin D’Haultfoeuille (2020) estimator

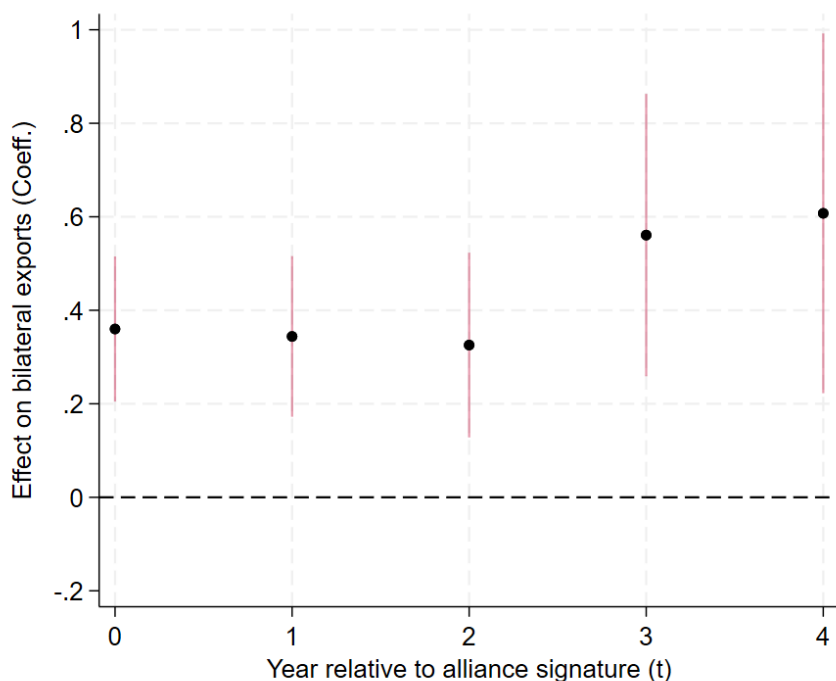
Dependent variable: Bilateral exports						
Time	t	t+1	t+2	t+3	t+4	t+5
Defence pact’s coeff	0.434 <sup>a</sup>	0.561 <sup>a</sup>	0.568 <sup>a</sup>	0.656 <sup>a</sup>	0.714 <sup>a</sup>	0.748 <sup>a</sup>
Standard error	(0.118)	(0.119)	(0.118)	(0.123)	(0.127)	(0.128)
RTA control	yes	yes	yes	yes	yes	yes
No. observ.	44.888	44.854	44.844	44.818	38.266	38.262
No. switchers	172	172	172	172	170	170

*Note:* Dependent variable is the inverse hyperbolic sine transformation of exports from country i to country j at time t in millions of current dollars. t is the year of the pair’s alliance signature. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

## 1.E The Cold War

How to define the Cold War and to what extent it can bias our estimation is not obvious. If we consider that the Cold War is a period structuring global country relationships, it is a time-level variable and is captured by our country-year fixed effects. If we define the Cold War as a period when countries were

Figure 1.D.1: Two-way robust fixed effects estimation, PPML



*Note:* Figure plots results of the two-way robust fixed effects estimation with PPML estimator and 95% confidence intervals. Dependant variable is exports.  $t$  is the year of the pair's alliance signature. Standard errors are clustered at the country-pair level.

either capitalist or communist, it comes under the country-year level and once again is captured by our fixed effects. Yet, if we consider the Cold War as a latent conflict between the western block and the eastern block, our fixed effects are not sufficient. Therefore, we include in our estimation a dummy variable taking the value one if country  $i$  is a member of one block and country  $j$  a member of the opposite block, conditional on the absence of military alliances between  $i$  and  $j$ , and zero otherwise.<sup>89</sup>

In this case, we deal with highly specific heterogeneity in alliance effects. The Cold War variable is a particular case of a non-alliance relationship between countries  $i$  and  $j$ . In section 1.3.3, the defence pact coefficient is estimated compared with the average case of non-alliances. Yet, introducing the Cold War variable would exclude a case of "latent-conflict-non-alliance" from this average. Therefore, the non-alliance average would be closer to a neutral relationship. This could induce overestimated coefficients. estimation should address this issue. Using propensity score matching, the treated and control groups should be comparable in terms of bilateral diplomatic relationships. Controlling for the  $Cold.war_{ijt}$  dummy, we estimate the defence pacts effect as in section 1.3.2 and with a DATT. The results are reported in table 1.E.1. In the standard estimation (column 1), controlling for the Cold War reduces the defence pact coefficient. In the DATT estimation (column 2), variables of interest

<sup>89</sup> $Cold.war_{ijt} = \mathbb{1}_{\{alliance_{ijt}=0\}} * ((alliance_{i,USA,t} + \mathbb{1}_{\{i=USA\}}) * (alliance_{RUS,j,t} + \mathbb{1}_{\{j=RUS\}}) + (alliance_{i,RUS,t} + \mathbb{1}_{\{i=RUS\}}) * (alliance_{USA,j,t} + \mathbb{1}_{\{j=USA\}})) * \mathbb{1}_{\{alliance_{i,USA,t} * alliance_{USA,j,t}=0\}} * \mathbb{1}_{\{alliance_{i,RUS,t} * alliance_{RUS,j,t}=0\}}$

coefficients are not impacted while the Cold War's coefficient is non-significant. In both estimations, the defence pact effect on bilateral exports is estimated at 80% and is highly significant.

Table 1.E.1: Defence pacts and the Cold War

Estimator: PPML		
Dependent variable: Exports		
Variables	Standard three-way FE	DATT
Defence pact	0.592 <sup>a</sup> (0.078)	
t-4		-0.035 (0.118)
t-3		0.065 (0.077)
t-2		-0.046 (0.086)
t-1		0.010 (0.088)
t		0.243 <sup>a</sup> (0.088)
t+1		0.380 <sup>a</sup> (0.085)
t+2		0.421 <sup>a</sup> (0.079)
t+3		0.518 <sup>a</sup> (0.078)
t+4		0.615 <sup>a</sup> (0.072)
>=t+5		0.575 <sup>a</sup> (0.043)
Weak Alliance	-0.127 (0.159)	-0.453 <sup>a</sup> (0.105)
Cold War	-0.239 <sup>a</sup> (0.076)	-0.029 (0.031)
RTA	0.135 <sup>a</sup> (0.032)	0.152 <sup>a</sup> (0.014)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	38.625

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; DATT, Differenced Average Treatment on the Treated; Dependent variable is exports from country  $i$  to country  $j$  at time  $t$ . In DATT, defence pact's signatory effect is estimated in comparison to  $k' \leq t - 5$ . Robust standard errors are in parentheses; column (1) standard errors are clustered at country-pair level; column (2) standard errors are clustered at country-pair-year level. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.



## 1.F Intensive and extensive margins of trade

In this section, we investigate through which margin of trade military alliances affect bilateral exports. We first estimate the effects of military alliances on bilateral exports conditional on positive flows from country  $i$  to  $j$  at year  $t$  and  $t-1$  (the intensive margin). Then, using a non-linear probability model (Kitazawa, 2012; Silva and Kemp, 2016), we estimate the effects of alliances on the probability of starting to export to a destination (the extensive margin).<sup>90</sup> Results are reported in appendix table 1.F.1. The intensive margin estimation results are reported in panel A of the table, while the extensive margin results are shown in panel B. First, alliances increase by 47% the bilateral exports of country-pairs that were already trade partners. Second, they increase by 35% country  $i$ 's probability of starting to export to country  $j$ . Thus, military alliances affect bilateral exports in terms of both margins. This suggests the presence of heterogeneous effects depending on which margin applies to the country-pair.

Table 1.F.1: Extensive and intensive margins of trade

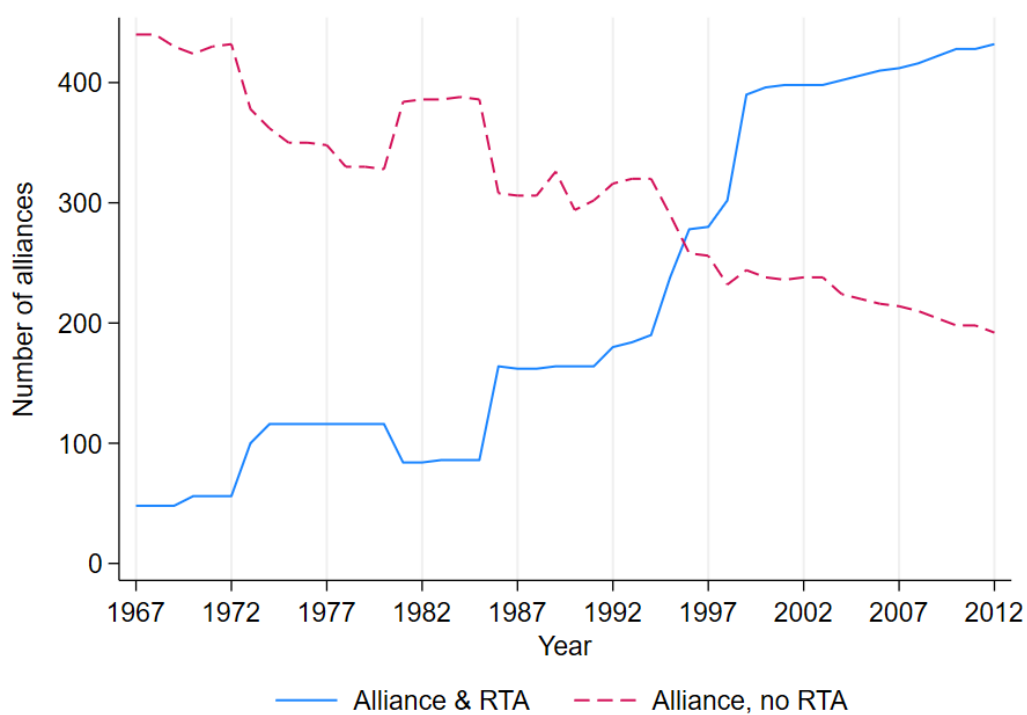
A: Intensive margin	
Dependent variable:	$\{X_{ijt} X_{ijt} > 0, X_{ijt-1} > 0\}$
Estimator:	PPML
Alliance	0.389 <sup>a</sup> (0.094)
Controls	yes
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	289.581
B: Extensive margin	
Dependent variable:	Export dummy
Estimator:	Logit
Alliance	0.403 <sup>a</sup> (0.079)
Controls	yes
Year FE	yes
Dyadic FE	yes
No. observ.	102.442

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country  $i$  to country  $j$  at time  $t$  without the zero observations and conditional to a positive value the previous year. Standard errors clustered at country-pair level are in parentheses. Intensive margin estimation's control: RTAs; Extensive margin estimation's controls: RTAs, exporter's GDP, importer's GDP, exporter's population, importer's population. Coefficient of our Logit estimation can be directly interpreted as the elasticity. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

<sup>90</sup>Given that the Logit estimator does not allow for intensive use of fixed effects, we use only country-pair and year fixed effects, which partially capture the multilateral resistance while controlling for importer/exporter GDP and population.

## 1.G Supplementary tables and figures

Figure 1.G.1: Alliances & RTAs, overlap



*Note:* The number of alliances counts each exporter-importer pair with an active alliance.

Table 1.G.1: List of countries

Albania	Denmark	Latvia	Saudi Arabia
Algeria	Ecuador	Kyrgyzstan	Russia
Argentina	Egypt	Libya	Serbia-Montenegro
Australia	Estonia	Lithuania	Singapore
Austria	Finland	Luxembourg	Slovakia
Bangladesh	France	Malaysia	Slovenia
Belarus	Gabon	Malta	South Korea
Belgium	Germany	Mexico	Spain
Bolivia	Greece	Morocco	Sri Lanka
Bosnia and Herzegovina	Hong Kong	Netherlands	Sweden
Brazil	Hungary	New Zealand	Switzerland
Brunei	Iceland	Nigeria	Taiwan
Bulgaria	India	North Macedonia	Thailand
Cameroon	Indonesia	Norway	Tunisia
Canada	Ireland	Pakistan	Turkey
Chile	Israel	Paraguay	Ukraine
China	Italy	Peru	United Kingdom
Columbia	Ivory Coast	Philippines	United States of America
Croatia	Japan	Poland	Uruguay
Cyprus	Kazakhstan	Portugal	Venezuela
Czech Republic	Kenya	Romania	Vietnam

Figure 1.G.2: Quantile estimates of RTAs

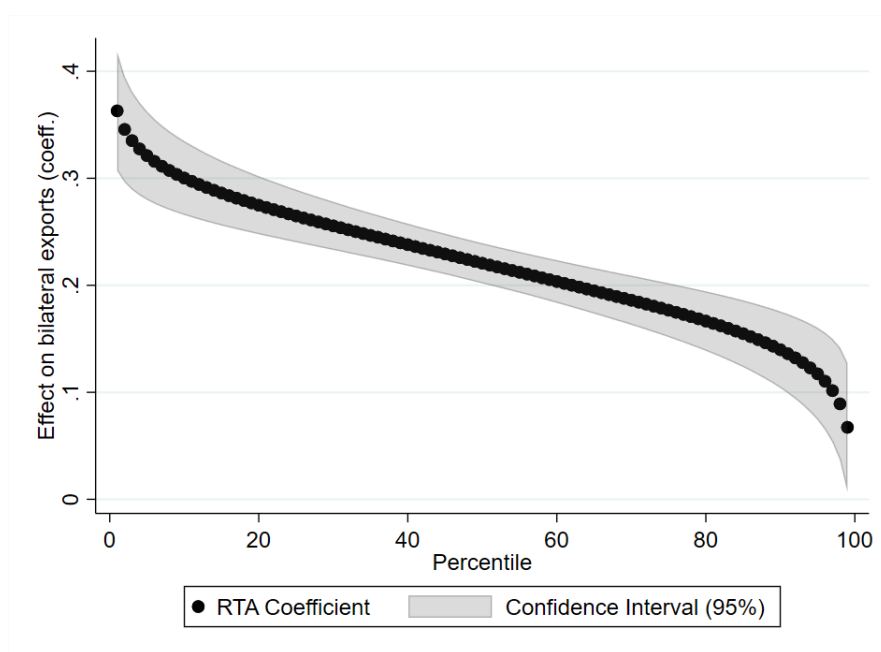


Table 1.G.2: Exports and military alliances, robustness checks

Robustness check:	Intra trade		Tariffs		Tariffs & Intra trade		RTA depth	Geo. and eco. dist.		Extended panel		Import shares	
Period:	1990-2012		1996-2012		1996-2012		1967-2012	1967-2012		1870-2012	1948-2012	1967-2020	1967-2012
Baseline sample:	No	Yes	No	No	No	No	No	Yes	No	No	No	No	Yes
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	
Alliance	0.490 <sup>a</sup> (0.046)	0.203 <sup>a</sup> (0.052)	0.232 <sup>a</sup> (0.059)	0.233 <sup>a</sup> (0.060)	0.353 <sup>a</sup> (0.043)	0.439 <sup>a</sup> (0.116)	0.438 <sup>a</sup> (0.116)	0.473 <sup>a</sup> (0.106)	0.554 <sup>a</sup> (0.065)	0.550 <sup>a</sup> (0.069)	0.366 <sup>a</sup> (0.082)	0.646 <sup>a</sup> (0.105)	
RTA	0.104 <sup>a</sup> (0.041)						-0.021 (0.052)	0.157 <sup>a</sup> (0.033)		0.165 <sup>a</sup> (0.030)	0.107 <sup>a</sup> (0.036)	0.374 <sup>a</sup> (0.045)	
ln(Tariffs+1)				-1.025 <sup>a</sup> (0.288)	-1.330 <sup>a</sup> (0.390)								
Intra x Year				-0.039 <sup>a</sup> (0.003)	-0.033 <sup>a</sup> (0.004)								
RTA depth						0.010 <sup>a</sup> (0.001)	0.011 <sup>a</sup> (0.002)						
Dif. GDP per cap.								-0.005 (0.004)					
ln(Distance) x Year								-0.000 (0.001)					
Exporter x Year FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	
Importer x Year FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	
Dyadic FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	
No. observ.	97,129	118,439	71,456	71,456	47,921	314,128	314,128	320,666	1,159,418	1,102,828	376,434	320,666	

Note: PPAI, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country i to country j at time t. Standard errors clustered at the country-pair level are in parentheses. Because of data limitation, several estimations are performed with a shorter panel or omitted observations/country-pairs. In column 3 we perform the same regression as in column 2 but without observations whose bilateral average tariff is unknown. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.3: Corrected three-way ppml estimation

Estimator: PPML	
Dependent variable: exports	
Variables	(1)
Alliance	0.482 <sup>a</sup>
Corrected bias	-0.010 (0.118)
RTA	0.166 <sup>a</sup>
Corrected bias	-0.005 (0.036)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	320.666

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. Coefficients are corrected from the asymptotic bias. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.4: Exports and defence pacts

Estimator:		PPML	
Dependent variable:		Bilateral exports	
Variables	(1)	(2)	
Defence pact	0.693 <sup>a</sup> (0.079)	0.694 <sup>a</sup> (0.120)	
Weak alliance	-0.107 (0.151)	-0.149 (0.153)	
RTA	0.137 <sup>a</sup> (0.032)	0.137 <sup>a</sup> (0.032)	
NATO		0.643 <sup>a</sup> (0.080)	
Exporter x Year FE	yes	yes	
Importer x Year FE	yes	yes	
Dyadic FE	yes	yes	
No. observ.	320.666	320.666	

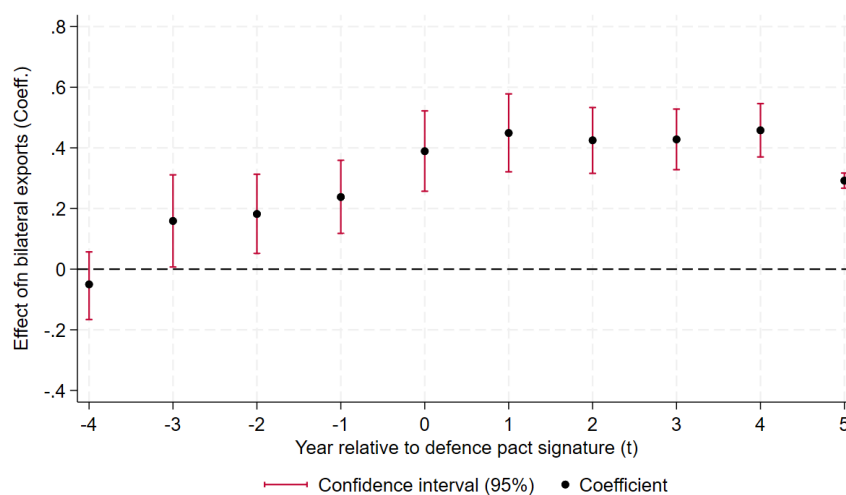
*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.5: Alliances and bilateral exports, 2SLS

Dependent variable: exports		
Second stage		
	(1)	(2)
Estimator:	OLS	2SLS
Instrument variable:	None	Common out. alliances
Alliance	0.664 <sup>a</sup> (0.088)	0.452 <sup>a</sup> (0.165)
RTA	0.562 <sup>a</sup> (0.029)	0.575 <sup>a</sup> (0.074)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.712	320.712
First-stage		
Instrumented variable:	None	Alliance
Common out. alliances		0.056 <sup>a</sup> (0.007)
RTA		0.012 <sup>a</sup> (0.008)
KPW rk F-stat:		67
KPW rk LM-stat:		11

*Note:* OLS, Ordinary Least Squares; 2SLS, Two-Stage Least Square; FE, Fixed effects; Dependent variable is the inverse hyperbolic sine transformation of exports from country  $i$  to country  $j$  at time  $t$  in millions of current dollars. Standard errors clustered at country-pair levels (column 1) and at the importer and exporter levels (column 2) are in parentheses. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Figure 1.G.3: Event study



*Note:* Figure plots results of the standard event study estimation of the defence pacts' effect on bilateral trade. Every time  $k$  is compared to  $k' \leq t_{def} - 5$ . The last period includes  $k \geq t_{def} + 5$ . Standard errors are clustered at the country-pair-year level.

Table 1.G.6: Alliances and bilateral exports, IV control function

Dependent variable: exports		
Second stage	(1)	(2)
Estimator:	PPML	PPML
Alliance	0.635 <sup>a</sup> (0.025)	
Defence pact		0.624 <sup>a</sup> (0.025)
Weak alliance		-0.011 <sup>a</sup> (0.041)
RTA	0.152 <sup>a</sup> (0.012)	0.144 <sup>a</sup> (0.011)
First stage residuals	-0.357 <sup>a</sup> (0.049)	0.176 <sup>a</sup> (0.057)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666
First-stage		
Estimator:	OLS	OLS
Dependent variable:	Alliance	Defence pact
Common out. alliances	0.056 <sup>a</sup> (0.007)	0.053 <sup>a</sup> (0.006)
Weak alliance		-0.398 <sup>a</sup> (0.062)
RTA	0.011 (0.008)	0.015 <sup>c</sup> (0.009)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.712	320.712
KPW F-stat	67	78
KPW LM-stat	11	10

*Note:* OLS, Ordinary Least Squares; FE, Fixed effects. Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. Standard errors clustered at the exporter and importer levels are in parentheses. Second-stage standard errors are bootstrapped. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.7: Outside alliances

Estimator: PPML		
Dependent variable: Exports		
Variables	(1)	(2)
Alliance	0.467 <sup>a</sup>	
	(0.108)	
Outside alliance(exp.)	-0.030	
	(0.174)	
Outside alliance(imp.)	-0.106	
	(0.157)	
Defence pact		0.702 <sup>a</sup>
		(0.081)
Weak alliance		-0.113
		(0.169)
Outside defence pact(exp.)		0.101
		(0.152)
Outside defence pact(imp.)		0.044
		(0.144)
Outside weak alliance(exp.)		-0.102
		(0.185)
Outside weak alliance(imp.)		0.101
		(0.266)
RTA	0.160 <sup>a</sup>	0.131 <sup>a</sup>
	(0.033)	(0.032)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	320.666	320.666

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. Dependent variable is exports from country *i* to country *j* at time *t*. Standard errors clustered at country-pair level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.



Table 1.G.8: Dynamic propensity score matching, some details

Estimator: Probit			
Dependent variable: Defence pact			
Variables	(1)	(2)	(3)
Year:	1967	1990	2012
Exp. ln(GDP)	0.379 <sup>a</sup> (0.031)	0.314 <sup>a</sup> (0.025)	0.318 <sup>a</sup> (0.027)
Imp. ln(GDP)	0.379 <sup>a</sup> (0.031)	0.314 <sup>a</sup> (0.025)	0.318 <sup>a</sup> (0.027)
Exp. ln(Pop.)	-0.185 <sup>a</sup> (0.035)	-0.117 <sup>a</sup> (0.029)	-0.116 <sup>a</sup> (0.028)
Imp. ln(Pop.)	-0.185 <sup>a</sup> (0.035)	-0.117 <sup>a</sup> (0.029)	-0.116 <sup>a</sup> (0.028)
Common religion	1.231 <sup>a</sup> (0.098)	1.616 <sup>a</sup> (0.098)	1.378 <sup>a</sup> (0.091)
ln(Distance)	-0.435 <sup>a</sup> (0.031)	-0.331 <sup>a</sup> (0.033)	-0.503 <sup>a</sup> (0.030)
Common official language	1.169 <sup>a</sup> (0.078)	1.023 <sup>a</sup> (0.082)	0.817 <sup>a</sup> (0.079)
Colonial past	-0.479 <sup>a</sup> (0.140)	-0.641 <sup>a</sup> (0.154)	-0.366 <sup>a</sup> (0.134)
No. observ.	6.970	6.870	6.866

*Note:* OLS, Ordinary Least Square; Dependent variable is the presence of a military alliance between countries *i* and *j* at time *t*. Standard errors are in parentheses. The PSM is made for each year from 1967 to 2012, respectively. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.9: Propensity score matching, variables' mean

Variable	Psm base	Standard base
GDP exp. (ln)	11.633	10.608
GDP imp. (ln)	11.629	10.608
Population exp. (ln)	2.816	2.575
Population imp. (ln)	2.812	2.575
Distance (ln)	7.762	8.511

Table 1.G.10: DATT

Estimator: PPML	
Dependent variable: Bilateral exports	
Variables	(1)
t-4	0.024 (0.117)
t-3	0.076 (0.075)
t-2	-0.035 (0.085)
t-1	0.021 (0.086)
t	0.254 <sup>a</sup> (0.087)
t+1	0.391 <sup>a</sup> (0.083)
t+2	0.431 <sup>a</sup> (0.078)
t+3	0.529 <sup>a</sup> (0.076)
t+4	0.626 <sup>a</sup> (0.070)
>=t+5	0.585 <sup>a</sup> (0.040)
Weak Alliance	-0.445 <sup>a</sup> (0.108)
RTA	0.152 <sup>a</sup> (0.014)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	yes
No. observ.	38.625

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. *t* is the date of the signatory of the alliance between country *i* and *j*. Defence pact's signatory effect is estimated in comparison to  $k' \leq t - 5$ . Standard errors clustered at country-pair-year level are in parentheses. Observations are weighted in function of our Propensity Score Matching. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.11: Defence pacts and UN votes

Estimator:	PPML	
Dependent variable:	Common votes (sum)	
Variables	(1)	(2)
Defence pact	0.020 <sup>a</sup> (0.004)	
Common out. def. pacts		0.001 <sup>a</sup> (0.000)
Weak alliance	0.009 (0.006)	0.007 (0.006)
RTA	0.033 <sup>a</sup> (0.002)	0.033 <sup>a</sup> (0.002)
Abstention	0.168 <sup>a</sup> (0.003)	0.168 <sup>a</sup> (0.003)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	251.806	251.806

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is the sum of common UN votes between country  $i$  and  $j$ . Abstention is the sum of the common abstention between country  $i$  and  $j$ . We use the inverse hyperbolic sine transformation of Abstention. For reasons of consistency with the baseline estimation, in structural gravity, each country-pair is counted twice ( $ij$  and  $ji$ ). Standard errors clustered at country-pair level are in parentheses. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.12: The military cooperation channel

Second stage	(1)	(2)
Estimator:	PPML	
Dependent variable:	Exports	
Military cooperation	0.401 <sup>a</sup> (0.069)	0.294 <sup>a</sup> (0.044)
Weak alliance	-0.227 (0.157)	-0.286 (0.178)
RTA	0.040 (0.036)	0.041 (0.040)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.304	167.304
First-stage	OLS	
Estimator:	OLS	
Dependent variable:	Military cooperation	
Defence pact	1.004 <sup>a</sup> (0.100)	
Common out. def. pacts		0.081 <sup>a</sup> (0.005)
Weak alliances	0.241 <sup>b</sup> (0.097)	0.178 <sup>c</sup> (0.096)
RTA	0.128 <sup>a</sup> (0.035)	0.125 <sup>a</sup> (0.035)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Dyadic FE	yes	yes
No. observ.	167.328	167.328

*Note:* OLS, Ordinary Least Square; PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects. The panel starts in 1989. Common defence pacts sum all external partners for which country *i* and *j* both have a defence pact. Military cooperation is the inverse hyperbolic sine transformation of  $\sum coop.mil.ev.ijt$ . Robust standard errors clustered at country-pair level are in parentheses. Second-stage standard errors are bootstrapped. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.13: Estimating the trade elasticity

Estimator: PPML	
Dependent variable: Bilateral exports	
Variables	(1)
ln(1+Tariffs)	-3.754 <sup>a</sup> (1.102)
ln(Distance)	-0.781 <sup>a</sup> (0.044)
Common religion	-0.157 (0.160)
Contiguity	0.579 <sup>a</sup> (0.111)
Common language	0.131 (0.101)
Exporter x Year FE	yes
Importer x Year FE	yes
Dyadic FE	no
No. observ.	54,479

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country *i* to country *j* at time *t* in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 1.G.14: GE exports and welfare gains from alliances (2012) with heterogeneous elasticities

Country (Iso3)	Exports	Real revenues	Country (Iso3)	Exports	Real revenues
ALB	2.53	0.20	ISL	26.34	9.52
ARG	13.55	4.07	ISR	-1.21	-1.04
AUS	2.64	0.26	ITA	20.32	2.85
AUT	-0.41	-1.25	JPN	5.10	0.32
BEL	11.47	11.80	KEN	-2.90	-0.17
BGD	0.32	-0.55	KOR	3.34	0.40
BGR	7.45	1.25	LKA	-0.77	-0.48
BOL	22.29	6.68	MAR	1.19	-0.15
BRA	12.81	0.57	MEX	16.47	7.03
CAN	24.08	6.14	MLT	1.01	-1.30
CHE	-0.16	-1.30	MYS	0.65	-0.49
CHL	8.99	3.93	NGA	1.30	-1.68
CHN	-0.33	-0.11	NLD	10.41	12.05
CIV	10.91	0.21	NOR	12.95	4.32
CMR	0.77	-0.54	NZL	-0.07	-0.26
COL	23.47	1.89	PAK	16.31	3.20
CYP	-0.33	-1.20	PER	17.16	1.70
DEU	15.54	3.87	PHL	3.21	1.07
DNK	17.04	6.10	POL	34.14	3.22
DZA	0.40	-0.50	PRT	33.81	4.53
ECU	21.34	3.90	ROM	2.66	0.04
EGY	-1.85	-0.38	SAU	0.58	-0.60
ESP	25.75	3.11	SER	-3.29	-0.27
FIN	-1.11	-0.78	SGP	-0.04	-0.24
FRA	24.91	4.31	SWE	-0.97	-1.47
GAB	1.31	3.06	THA	0.42	-0.44
GBR	22.94	4.09	TUN	1.56	-0.22
GRC	36.39	3.59	TUR	35.42	1.75
HKG	0.28	-0.71	URY	29.96	2.76
HUN	17.40	7.04	USA	34.09	2.02
IDN	0.32	-0.21	VEN	11.75	5.72
IND	-1.13	-0.15	VNM	0.36	-0.29
IRL	1.35	-2.07			

*Note:* The real revenue is our measure of welfare. All numbers are variations in percentage.

Table 1.G.15: Alliances scenarios and welfare changes

Scenarios:		NATO expansion		NATO East-rupture		New East-block		Scenarios:		NATO expansion		NATO rupture		New East block	
Country (Iso3)	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues	Country (Iso3)	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues	Real revenues
ALB	23.81	-1.24	6.24	ISL	1.60	-0.37	-0.45								
ARG	-0.20	0.04	0.01	ISR	-0.69	0.12	0.12								
AUS	3.29	0.02	0.02	ITA	1.08	-0.27	-0.33								
AUT	17.46	0.40	0.20	JPN	1.61	0.01	-0.00								
BEL	1.80	-0.38	-0.46	KEN	-0.17	0.01	-0.02								
BGD	-0.49	0.07	-0.08	KOR	3.07	0.03	0.01								
BGR	10.27	-2.67	2.52	LKA	-0.39	0.03	-0.03								
BOL	-0.15	0.00	-0.02	MAR	-0.39	0.09	0.05								
BRA	-0.07	0.01	0.01	MEX	-0.25	0.04	-0.01								
CAN	1.28	-0.02	-0.03	MLT	23.44	0.25	0.20								
CHE	17.13	0.19	0.15	MYS	-0.54	0.04	0.00								
CHL	-0.46	0.06	0.02	NGA	-1.03	0.12	0.06								
CHN	-0.07	0.01	0.09	NLD	2.27	-0.50	-0.65								
CIV	-0.06	0.02	0.01	NOR	2.63	-0.23	-0.42								
CMR	-0.38	0.06	0.03	NZL	6.47	0.02	0.00								
COL	2.05	0.02	0.00	PAK	6.83	0.07	0.81								
CYP	-1.24	0.32	0.25	PER	-0.24	0.01	-0.00								
DEU	1.88	-0.59	-0.67	PHL	-0.46	0.03	0.02								
DNK	3.40	-0.46	-0.72	POL	0.73	-5.60	-3.50								
DZA	-0.38	0.09	0.07	PRT	0.57	-0.10	-0.12								
ECU	-0.26	0.03	-0.01	ROM	11.20	-0.65	3.98								
EGY	-0.33	0.07	0.02	SAU	-0.71	0.08	0.09								
ESP	0.68	-0.19	-0.22	SER	6.52	0.24	3.25								
FIN	11.51	0.13	4.43	SGP	-0.27	0.01	0.02								
FRA	1.00	-0.22	-0.25	SWE	15.73	0.19	3.38								
GAB	-0.82	0.16	0.29	THA	-0.54	0.04	0.01								
GBR	1.71	-0.22	-0.26	TUN	-0.53	0.12	0.09								
GRC	0.96	-1.04	-1.25	TUR	0.40	-2.73	-1.64								
HKG	-1.14	0.07	0.18	URY	-0.15	0.03	0.01								
HUN	2.93	-12.05	-5.57	USA	0.03	-0.03	-0.04								
IDN	-0.32	0.02	-0.00	VEN	-0.31	0.03	0.03								
IND	-0.13	0.02	0.00	VNM	-0.33	0.03	-0.03								
IRL	24.58	0.18	0.11												

Note: The real revenue is our measure of welfare. All numbers are variations in percentage.

## Chapter 2

# Firms under fire!

## How insecurity affects formal firms' existence

This chapter studies the effect of insecurity on formal firms' existence. We develop a flexible theoretical framework in which insecurity affects firms' market entry, exit, and formality decisions. In our empirical analysis, we combine an original dataset on Afghan firms with georeferenced data on military events during the post-2003 Afghan conflict. In such a state-building context, exposure to military events has an average positive effect on formal firms' existence. Nonetheless, this effect is highly heterogeneous depending on actors, location, timing and firms' characteristics. The Afghan conflict has the specificity of deeply involving foreign countries. Mobilising this particular source of exogenous variation, we identify insecurity's causal effect on formal firms' existence. We show that an increase of 1% in the exposure to instrumented military events raises the formal activity probability by 4.16%.

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The present chapter studies the effect of insecurity on formal firms' existence. Insecurity is generally defined as a production or exchange cost induced by exposure to violence ([Anderson and Marcouiller, 2002](#); [Martin et al., 2008](#); [Couttenier et al., 2023](#); [Del Prete et al., 2023](#)): goods and production factors can be hijacked, expropriated, or simply destroyed, while infrastructure or, more generally, market access can be severely damaged. Its impact on formal firms is a primary concern. Formal economic activities are vital for economic and institutional development. They concentrate the most productive firms and are indispensable to states' finances ([De Paula and Scheinkman, 2011](#); [Ordonez, 2014](#); [Ulyssea, 2018](#)).

To analyse this microeconomic impact of insecurity, conflict events<sup>1</sup> are commonly used as a direct proxy ([Verwimp et al., 2019](#)). Yet, the literature's focus is mainly context-specific and on short-term effects, with the common primary assumption that all conflict events increase the insecurity of all economic agents. Investigating the consequences of the 2008 Georgia-Russia war, [Petracco and Schweiger \(2012\)](#) show that temporary exposure to conflict reduced firms' next-year activity. In the same way, [Blumenstock et al. \(2018\)](#) study the impact of Afghan districts' exposure to major terrorist attacks on firms. Using phone data on corporate line customers, the paper concludes that major terrorist events reduce firms' activity over the following months. Exploiting spatial variations in the 2000 Cote d'Ivoire civil conflict, [Klapper et al. \(2013\)](#) analyse the negative consequences on firms' productivity. [Camacho and Rodriguez \(2013\)](#) show that an increase in Columbian firms' municipality exposure to guerilla and paramilitary attacks increases the probability of firm exit. Interestingly, they also highlight that the effect of insecurity is stronger on small firms. Focusing on the 2nd Libyan civil-war (2014-2020), [Del Prete et al. \(2023\)](#) demonstrate that exposure to conflict has a negative non-linear effect on firms' output. Developing a production network analysis, [Couttenier et al. \(2022\)](#) estimate the direct and indirect effects of the Maoist insurgency on Indian firms. Their results show an average aggregate output loss of 3.8 billion USD per year.

Yet, conflicts are very heterogeneous (actors, objectives, duration, etc...). The average effects of exposure to conflict may hide a large diversity of treatments and consequences. Especially, in certain cases, conflict events may result, in the long-run, in an improvement of security through state-building. The provision of security, i.e. peace building, is the first goal of state-building. Conflict events may capture the strengthening of the state over its territory, and, therefore, signalling an expansion of the state capacity, convey a reduction in insecurity ([Berdal, 2017](#); [Fritz and Menocal, 2007](#)).

The present chapter addresses this issue. Using detailed data on the Afghan firms and conflict events, we show that, in a specific context, exposure to conflict events has in the long-run an average

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<sup>1</sup>Conflict events are incidents where armed force was used by an organised actor against another organized actor, or against civilians, resulting in at least 1 direct death at a specific location and a specific date ([Sundberg and Melander, 2013](#)).

positive effect on formal firms' existence, but with a large heterogeneity depending on actors, location, timing and firms' characteristics. Mobilizing exogenous variations in conflict events, we develop a method to analyse the causal effect of insecurity on formal firms' existence despite the complexity and diversity of conflicts. The post-2003 Afghan conflict provides a remarkable context. The existence of the formal market (and formal firms) requires the presence of a state claiming for the territory. In an unstable country such as Afghanistan, this condition may be discussed. Nonetheless, since 2003 we have been able to clearly identify an Afghan state. Before, identifying a state in Afghanistan is difficult. Even if the Taliban controlled Kabul, they were unable to establish a central power, had no robust internal legitimacy or administration, and were not recognized on the international scene. In 2003, the Taliban had already lost Kabul, and the local actors opposed to the Taliban, including the transitional government and the warlords, agreed on the necessity of a strong central state. In 2004 the Loya Jirga (the legislative assembly) enacted a constitution with strong presidential power. A few months later, elections were organized with good support from the population. Thus, since 2003 there has been an institution able to exercise the state functions and recognized by the other nations (Barfield, 2010; Maley, 2020).<sup>2</sup>

Yet, this state struggles in a military competition for a territory with a large diversity of actors whose objectives and strategies have been intensively communicated and documented (Fetzer et al., 2021). In a general environment of instability, two main camps – the Taliban versus the Afghan state and its allies from the International Security Assistance Force (ISAF)<sup>3</sup> – fight each other for the control of the country. This confrontation is materialised by a collection of military events – conflict events opposing one or several states to another organized actor.<sup>4</sup> In such a context of state-building with competition for territory control, military events should capture the investment from the state to bring security and territory control, signal an increase in state capacity and positively affect formal firms (Rocha Menocal, 2011; Thies, 2007). Yet, if driven by the Taliban's strengthening, military events may, on the contrary, capture a direct threat. Therefore, this context presents the perfect conditions to dive into the complexity of the relationship between conflict events, insecurity, and formal firms' activity.

We empirically investigate this question. Our database combines conflict events observations from the geocoded Uppsala Conflict Data Program (UCDP) project with a set of Afghan firms from the former Afghan Ministry of the Economy register of formal companies. We end with a panel of 37308

<sup>2</sup>The taking of Kabul by the Taliban and the dissolution of the republic questions again the presence of state in Afghanistan. However, our study being limited to the period 2003-2016, this does not threaten the identification of an Afghan state.

<sup>3</sup>ISAF is an initiative from the Northern Atlantic Treaty Organisation (NATO). Yet, some countries that are not members of NATO participate in ISAF, such as Sweden or South Korea.

<sup>4</sup>Following UCDP definition, and organized actor is a state or a non-state formally organised or informally organised group. Military events, therefore, exclude non-state violence – events where none actor is a state – and one-sided violence – events where civilians are directly targeted.

firms from 2003 to 2016. The Afghan conflict is characterized by the confrontation between the state coalition (Afghan state & ISAF) and the Taliban through military events. Over the period, 20416 military events are reported. As we use georeferenced data, both firms and events are localised over 200 districts. We develop a theoretical model where heterogeneous firms endogenously enter the market, exit, and choose their (in)formality status. In this framework, we introduce insecurity – a marginal cost capturing the production losses due to destruction and expropriation. We show how this affects firms' choices and derive the insecurity elasticity in the formal activity probability. The model predicts that a reduction of 1% in insecurity increases the formal firms' existence between 7.86% and 11.86%.

Based on the theoretical expression of the formal activity probability, we show that, on average, exposure to military events has a small positive effect on formal firms' existence. In the 2003-2016 Afghan state-building context, military events can entail important damages but mainly signal an expansion of the state capacity, reduce formal firms' insecurity and, therefore, increase the probability of formal activity. Digging into the heterogeneity of military events, we show their effect highly depends on actors, timing, and location.

The actor initiating a military event may strongly matter. For formal firms, states are associated with a reduction in insecurity and the Taliban with a rise. Yet, this information is not directly available in data. To address this issue as well as any endogeneity concerns, we develop an Instrumental Variable (IV) strategy. We follow a [Bartik \(1991\)](#) *shift-share* approach. The Afghan conflict has the particularity of strongly involving foreign countries. Relying on this specificity, we estimate the impact of a change in military events determined by the geographic allocation (share) of exogenous variations in NATO's military events (shift). By doing so, we constrain military events to vary depending on a unique belligerent and can identify a clear change in insecurity. Thus, an increase of 1% in the exposure to instrumented military events raises the formal activity probability by 4.16%. Alternative IVs and a plausible exogeneity test based on [Conley et al. \(2012\)](#) confirm the robustness of our results. Using the richness of our database, we test the presence of heterogeneous effects depending on firms' characteristics. Although all formal firms benefit from district-level reductions in insecurity, small firms and those with a NATO member nationality are favoured, which strongly distorts the supply structure.

The chapter is organised as follows. Section 1 presents the theoretical model. Section 2 describes our dataset and some descriptive statistics. Section 3 analyses the effect of exposure to military events on Afghan firms' formal activity probability. Section 4 details our strategy to identify the exogenous effect of an insecurity reduction. In the end, we bring a short conclusion in section 5.

## 2.1 Formal activity probability and insecurity costs

This first section presents the theoretical model. In this framework, we develop how insecurity affects the conditions a firm must fulfil to be active in the formal market and the probabilities associated. It combines heterogeneous firms with endogenous market entry, exit, and (in)formality status.

### 2.1.1 Set-up

Consumers maximise their utility by consuming  $q(\omega)$  units of each differentiated good  $\omega$ , so that  $U = [\int_{\Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} d\omega]^{\frac{\sigma}{\sigma-1}}$ . The output of a given firm is defined by the linear production function  $q = \gamma l - f$ , with  $\gamma$  the firm's productivity,  $l$  labour and  $f$  a fixed cost. Two sectors coexist: the formal sector and the informal sector. Firms cannot produce in both sectors. They chose to be fully formal or informal. They produce the same variety but with different production functions depending on their choice. Following [Ulyssea \(2018\)](#), the informal sector is characterized by a marginal cost distortion: we consider  $\gamma$  and  $\gamma_{inf}$ , respectively the firm's productivity level in the formal and in the informal sectors.

In keeping with [Chaney \(2008\)](#) and [Helpman et al. \(2008\)](#), we assume the distribution of firms' productivity follows a Pareto distribution. Each sector follows its own Pareto distribution of productivity.<sup>5</sup> In the formal sector, productivities follow the distribution  $G(\alpha) = \frac{\alpha^\theta}{\bar{\alpha}^\theta}$ , while in the informal sector productivities are distributed following  $G_{inf}(\alpha_{inf}) = \frac{\alpha_{inf}^\delta}{\bar{\alpha}_{inf}^\delta}$  where  $\alpha = \frac{1}{\gamma}$ ,  $\bar{\alpha}$  is the maximum marginal cost, and  $\theta$  and  $\delta$  are Pareto shape parameters.<sup>6</sup> To match with the stylized fact of smaller and more uniform informal firms ([De Paula and Scheinkman, 2011](#); [Meghir et al., 2015](#); [Ordonez, 2014](#); [Pratap and Quintin, 2006](#)), we set the distortion  $\delta > \theta$ <sup>7</sup> and  $\bar{\alpha}_{inf} < \bar{\alpha}$ .<sup>8</sup> This distortion on the technological parameters  $\delta$  and  $\alpha_{inf}$  implies that, on the informal market, firms are less likely to observe high productivity, but are less penalized by a low one. In other words, this productivity distortion can be interpreted as better suitability of the informal production organization for unproductive firms than for the highly productive ones.<sup>9</sup>

In standard heterogeneous firms models, firms always expect the right costs at equilibrium. Therefore, all exits are purely exogenous ([Melitz, 2003](#)) and *inefficient* creations are not possible. We relax this assumption by integrating imperfect information on fixed costs  $f$ . Firms face the following antic-

<sup>5</sup>For solving issues we follow [Chaney \(2008\)](#) assuming the minimum marginal cost is equal to 0.

<sup>6</sup>For simplicity, here we assume that firms independently draw their formal and informal productivities. Therefore, it allows firms to draw productivities that place them differently in the formal and informal productivity distributions. This assumption affects parameters' elasticities in the final expression of the formal activity probability. In section [2.B.2](#), we further discuss this point and extend the model with an informality structure where firms place on the productivity distributions is restricted.

<sup>7</sup>This means that the distribution of firms' productivity is more concentrated in the informal market.

<sup>8</sup>The minimum productivity firms can draw in the informal market is higher than on the formal market. This facilitates firms entry in the informal market.

<sup>9</sup>Indirectly, the productivity distortion also captures the structural risk of detection and sanction in the informal sector ([Ulyssea, 2018](#)).

ipation problem. Once on the market, firms draw a fixed cost following an inverse Pareto distribution  $P(F < f^*) = \frac{f^{*k}}{\bar{f}^k}$ , with  $0 < k < 1$  and  $\bar{f}$  the maximum fixed cost in the economy. The shape of the fixed costs' distribution is determined by  $k$ . The lower  $k$ , the more the distribution is concentrated around zero. If the drawn fixed cost  $f^*$  is larger than the expected fixed cost  $\underline{f}$ , the firm must pay  $f^*$ , otherwise it pays  $\underline{f}$ . Therefore, the large majority of firms face  $\underline{f}$ , but few observe a higher fixed cost. As we detail below, aleatory fixed costs allow for endogenous exits.

### 2.1.2 Firm's probability of formal existence

We can derive the firm's probability of formal existence. We assume firms based their decisions on their expected profit.<sup>10</sup> A firm exists on the formal market under three conditions. First, it must enter the market. A firm enters if its expected formal profit is superior to zero:

$$\frac{E}{P^{1-\sigma}} \frac{(\frac{\sigma}{\sigma-1}w\alpha)^{1-\sigma}}{\sigma} - \underline{f} > 0 \quad (2.1)$$

where  $P$  is the price index,  $w$  the wage,  $E$  the total expenditures and  $\sigma$  the elasticity of substitution. Second, it must choose to be formal. In our frame, this choice is driven by the difference in the firm's formal and informal productivities – their distribution capturing (in)formality-specific production conditions. Therefore, knowing  $\gamma = \frac{1}{\alpha}$ , it must respect:<sup>11</sup>

$$\alpha < \alpha_{inf} \quad (2.2)$$

Third, it must not exit the market. Once the firm  $i$  knows its productivity ( $\gamma_i = \frac{1}{\alpha_i}$ ) and has decided to enter the market, it discovers its real fixed cost and must choose whether it stays or leaves the market: if a firm expected a lower fixed cost ( $f$ ) than the drawn one ( $f^*$ ), and is not sufficiently to conserve a positive profit, it exits. Thus, to remain on the market (as a formal firm), the firm must respect the condition:

$$\frac{E}{P^{1-\sigma}} \frac{(\frac{\sigma}{\sigma-1}w\alpha_i)^{1-\sigma}}{\sigma} - f^* > 0 \quad (2.3)$$

The productivity and the fixed cost are randomly distributed following their respective distributions  $G(\alpha)$  and  $P(f^*)$ . Therefore, from these three conditions we can derive  $\Gamma_e$  the probability of entry,  $\Gamma_{for}$  the probability of being formal, and  $\Gamma_x$  the probability of non-exit, detailed in the appendix section

<sup>10</sup>In such a context like Afghanistan, we may expect that entrepreneurs or firms also take their economic decision based on political benefits. For simplicity we do not mention them directly in the model. Yet, this can be easily done either by assuming that political benefits are directly transferable into economics returns for the firm, or by introducing (political) preferences of firms for (in)formality. This does not affect conclusions of the model.

<sup>11</sup>Here we chose the simplest modelling of the informal sector. Thus, all the formality probability relies on drawing, independently, a formal productivity higher than the informal one. Nonetheless, extensions of this modelling and a discussion of their consequences are provided in section 2.B.

2.A. Once combined, we obtain the firm's formal activity probability:<sup>12</sup>

$$\Gamma = \Gamma_e * \Gamma_{for} * \Gamma_x = \mu \left( \frac{E}{P^{1-\sigma}} \right)^{\frac{\theta}{\sigma-1} + k} \alpha_{inf}^{\theta} w^{-\theta - k(\sigma-1)} \bar{\alpha}^{-2\theta} \alpha^{k(1-\sigma)} \underline{f}^{-\frac{\theta}{\sigma-1}} \bar{f}^{-k} \quad (2.4)$$

Each parameter's elasticity depends on its implication in the three probabilities which compose  $\Gamma$ . They all follow the profit maximisation logic: the aggregated real expenditures ( $\frac{E}{P}$ ) and the informal marginal cost ( $\alpha_{inf}$ ) positively affect the formal activity probability, while all parameters related to the formal production costs ( $w; \bar{\alpha}; \alpha; \underline{f}; \bar{f}$ ) affect this probability negatively.

### 2.1.3 Insecurity costs

We model the insecurity cost as a variable cost  $\psi > 1$  (Anderson and Marcouiller, 2002; Martin et al., 2008; Couttenier et al., 2023; Del Prete et al., 2023). The insecurity cost is interpretable as a hidden production or exchange cost induced by exposure to violence (destruction, terrorism, conflicts, war, armed factions). When firms see a part of their production being expropriated or destroyed, the number of products that reach the consumer is lower – as if one unit of labour hired by the firm would no longer produce  $\gamma q$  but  $\frac{\gamma}{\psi} q$ .<sup>13</sup> Therefore, in an insecure environment, formal firms face the price function:  $p(\alpha) = \frac{\sigma}{\sigma-1} w \psi \alpha$ .<sup>14</sup>

Integrating insecurity costs in the model, we obtain the following insecurity-extended entry (2.5), (in)formality (2.6) and non-exit (2.7) conditions:

$$\frac{E}{P^{1-\sigma}} \frac{\left( \frac{\sigma}{\sigma-1} w \psi \alpha \right)^{1-\sigma}}{\sigma} - \underline{f} > 0 \quad (2.5)$$

$$\alpha \psi < \alpha_{inf} \quad (2.6)$$

$$\frac{E}{P^{1-\sigma}} \frac{\left( \frac{\sigma}{\sigma-1} w \psi \alpha_i \right)^{1-\sigma}}{\sigma} - f^* > 0 \quad (2.7)$$

and the insecurity-extended formal activity probability:

$$\Gamma = \mu \left( \frac{E}{P^{1-\sigma}} \right)^{\frac{\theta}{\sigma-1} + k} \alpha_{inf}^{\theta} \psi^{-2\theta - k(\sigma-1)} w^{-\theta - k(\sigma-1)} \bar{\alpha}^{-2\theta} \alpha^{-k(\sigma-1)} \underline{f}^{-\frac{\theta}{\sigma-1}} \bar{f}^{-k} \quad (2.8)$$

<sup>12</sup>with the mark-up constant  $\mu = (\sigma - 1)^{\theta + k(\sigma-1)} \sigma^{\theta \frac{\sigma}{\sigma-1} - k\sigma}$  derived from the mark-up's  $\frac{\sigma}{\sigma-1}$  and the elasticity of substitution's  $\frac{1}{\sigma}$  impacts on the entry and exit probabilities through the profit function.

<sup>13</sup>Alternatively, we may prefer to interpret it as a risk parameter. Because of the exposure to violence, firms fear losses in productivity, demand or more generally in expected profit. They internalise this risk in the price function. Lower insecurity means a lower risk and, therefore, lower prices.

<sup>14</sup> $\psi$  is directly interpretable as the probability of product to be expropriated (Anderson and Marcouiller, 2002; Martin et al., 2008). The uncertainty of insecurity applies at the product level. Firms are in an insecure environment but know to what extent they are subject to insecurity and how costly it is. While at the product level insecurity can be interpreted as a probability of being destroyed/stolen, at the firm level the expropriated share is known. Furthermore, it implies that firms facing the same level of insecurity will face the same insecurity cost  $\psi$ . Because there is this uncertainty at product level, the firm may want to insure each product against the risk that it is expropriated. Yet, it does not affect our model. If the firm insures its production, it must pay  $(1 - \psi) * 100$  percent of the product value in counterparts. Aggregating at the firm's total production level, it is equivalent to a variable cost  $\psi$ .

We can identify the insecurity costs' elasticity  $\varepsilon(\psi) = -2\theta - k(\sigma - 1)$ . It translates the sensitivity of each existence condition to insecurity (cf. eq 5, 6 & 7). Entry and informality conditions' sensitivities are derived from the shape of the productivity distribution  $\theta$  – i.e. from the chance of having sufficiently high productivity to have positive profit despite insecurity. Exit condition's sensitivity is, on the other hand, derived from the shape of the fixed cost distribution  $k$  and the elasticity of substitution  $\sigma$  – i.e. from the risk of expecting a fixed cost too low compared to the drawn one to maintain a positive profit. The insecurity elasticity is negative.<sup>15</sup> Higher insecurity costs reduce firms' formal profit, and thereby, the capacity to fulfil the existence conditions (cf. eq 5, 6 & 7). Therefore, any event that would increase the insecurity cost lowers the firms' probability of formal activity.<sup>16</sup>

#### 2.1.4 Model's predictions

Using this model, we can predict the reaction of firms to an insecurity shock. To do so, we need to calibrate parameters that compose the insecurity elasticity.  $\theta$ , the formal productivity distribution shape parameter, is the main component. It defines the sensitivity of the probability of entry and the probability of formality to insecurity.<sup>17</sup> In the literature, the standard value of  $\theta$  is 4 (Head et al., 2014). In the Afghan case, we estimate the value of  $\theta = 3.93$  (see appendix section 2.C). The second key parameter is  $\sigma$ , the elasticity of substitution. It defines the sensitivity to a change in insecurity of firms' output, revenue  $(1 - \sigma)$  and non-exit probability  $(-k(\sigma - 1))$ . Unfortunately, we cannot observe or estimate  $k$ , the shape parameter of the fixed costs distribution. Yet, we know from the theory that  $k$  is between 0 (i.e perfect information) and 1 (i.e no information).

Therefore, with  $\sigma = 5$  (Head et al., 2014) and  $\theta = 3.93$ , our model predicts that a reduction in insecurity of 1% would increase firms' output and revenue by 4% ( $\sigma - 1$ ), raises the probability of entry and the probability of formality by 3.93% ( $\theta$ ), and the probability of staying on the market between 0 and 4% ( $k(\sigma - 1)$ ). Finally, with the insecurity elasticity  $\varepsilon(\psi) = -2\theta - k(\sigma - 1) = -7.86 - 4k$  the model also predicts that a reduction in insecurity of 1% increases formal firms' existence between 7.86% and 11.86%. The insecurity elasticity's size is rather important. In theoretical frames with CES utility function and a Pareto distribution of productivities, firms' output, profit and entry are always very sensitive to marginal costs changes. The size of the insecurity elasticity also reflects the difficulties firms must face to formally exist through the combination of necessary conditions (entry, formality, non-exit).<sup>18</sup>

<sup>15</sup>by construction  $\sigma > 1$  (Chaney, 2008; Head and Mayer, 2014; Melitz, 2003)

<sup>16</sup>In our model, firms can choose to be formal or informal but formal and informal productivities, wages and insecurity costs are considered independent. In appendix section 2.B we propose extensions relaxing this assumption. Discussing the consequences on the insecurity elasticity, we show that stronger connections between formality and informality make firms more sensitive to insecurity. The standard insecurity costs' elasticity ( $\varepsilon(\psi) = -2\theta - k(\sigma - 1)$ ) must therefore be interpreted as a lower bound elasticity.

<sup>17</sup>The Entry, (in)formality and non-exit probabilities are detailed in the appendix section 2.A.

<sup>18</sup>For instance, in an economy with no possibility of informality and perfect information on fixed-costs the predicted

## 2.2 Data

So far, we discussed the theoretical effect of insecurity on formal firms' existence. Based on this model and using a large dataset combining precise information on Afghan firms and detailed observations of military events, we empirically investigate in the rest of the chapter the effect of insecurity on firms' formal activity probability.

### 2.2.1 Data description

*Firms.* – Data on Afghan firms are retrieved from the 2003-2016 register of formal companies of the Afghan Ministry of Economy. This register contains all firms declared to the state at least one year from 2003 to 2016 for a total of 37308 firms. Those firms legally exist. They benefit from access to the state's institutions and protection, but must comply with the law and pay taxes in exchange. Therefore, as long as they are registered, they are present on the formal market.

In our data, those declared firms and their characteristics are precisely identified. We know their location,<sup>19</sup> their number of workers at the year of entry, the name and nationality of the president, and their sector. Each firm has a specific ID. That allows us to follow them over the different years and categories of the register. For each of them, we observe the year of their entrance in the register and the year of exit. Once a firm exits, it cannot enter again. The entrance of a firm into the register is not automatic. A firm can be created, active but undeclared, and then become formal after a certain time if declared. In the same way, exit can be the consequence of full closure or a switch to informality. Therefore, based on this registering system, we can track firms' formal existence over years and districts.

*Military events.* – Based on firm geographic location, we merge each firm-year with information on military events from the geocoded Uppsala Conflict Data Program (UCDP) project (Sundberg and Melander, 2013). In the literature, military events are the gold standard to approximate insecurity changes.<sup>20</sup> Formally, military events are *incidents where armed force was used by an organised actor against another organised actor, whose at least one of them is a state, resulting in at least one direct death at a specific location and a specific date.*<sup>21</sup> For each event, we know the actors involved. We also know their location at the district level, and the year of their occurrence. With this information, we

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insecurity elasticity would be much lower (i.e.  $-\theta = -3.93$ )

<sup>19</sup>After harmonization work, we are able to determine the district of location of each firm. More precise location information is not exploitable for econometric analysis.

<sup>20</sup>For instance: Camacho and Rodriguez (2013); Couttenier et al. (2022); Del Prete et al. (2023); Martin et al. (2008).

<sup>21</sup>Even if in Afghanistan they are a minority, the UCDP database also contains some conflict events which oppose non-state actors. Yet, because these events (or their absence) may be the consequence of the military events' occurrence, including them may lead to biased estimates. Thus, in our empirical analysis, we use only military events and therefore focus on the conflict opposing states against the Taliban and Daesh for the control of the Afghan territory. The endogeneity of military events and their validity as a proxy for insecurity costs is discussed in detail in further sections.



can easily match our data on firms. We create the variable  $Mil.ev_{it}$ , a continuous variable summing all the military events the firm  $i$  was exposed from year 2003 to  $t$ . For interpretation reasons, we use in our empirical analysis the inverse hyperbolic sine transformation of this sum.<sup>22</sup> In this way, we put less weight on extreme observations, interpret coefficients as elasticities and keep the null observation that a simple log-linearisation would have dropped.<sup>23</sup>

## 2.2.2 Descriptive statistics

Our final database is a panel of 37308 formal firms from 2003 to 2016.<sup>24</sup> During our period, all observed firms enter the formal market for at least one year, but at different times. On the other hand, only 3.05% of them experience an exit.<sup>25</sup> Entries and exits are well dispersed over time.<sup>26</sup> Firms are split into 21 sectors. The sector categories are determined by the Afghan ministry of economy. Both industry and services are represented (see appendix table 2.D.2 for more details). On average firms have 20 workers. The distribution of firms' size in workers is very close to Pareto distribution. A large number of firms are very small, while few are enormous. 50% of firms have less than 6 workers, 75% less than 11, 90% less than 41 and 99% less than 201. The maximum number of workers observed is 11928.<sup>27</sup> The large majority of firm presidents<sup>28</sup> – 91.2% – are Afghan. Yet, we also observe 86 other nationalities.<sup>29</sup> In our sample, 5% of firm presidents are NATO member nationals, 3.6% are nationals of other foreign countries, and 0.2% have an unknown nationality. Figure 2.2.1.a presents the number of firms of our sample by district. Although Kabul concentrates a huge part of the formal activity, firms are well dispersed over 200 districts.

In Afghanistan, although industrial and service sectors represent more than half of the GDP, informality is endemic. In 2012, only about 30% of firms were declared while, excluding agriculture, formal labour represents less than 25% of total employment.<sup>30</sup> Nonetheless, these formal 30% constitute

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<sup>22</sup>Formally,  $Mil.ev_{it} = \ln \left[ \sum_{2003}^t (military.events_{it}) + \sqrt{\sum_{2003}^t (military.events_{it}) + 1} \right]$ .

<sup>23</sup>Knowing that on average, before transformation,  $Mil.ev_{it} = 82 > 10$ , we are not exposed to approximation errors (Bellemare and Wichman, 2020).

<sup>24</sup>Even when the firm does not exist on the formal market, it is considered as observed. Thus, we assume that each economic area has a potential stock of firms. Some of them exist on the formal market, some are informal, and certain are "sleeping". In other words, even if the firm does not exist on the formal market, entrepreneurs do and are exposed to insecurity.

<sup>25</sup>Exiting firms' characteristics are similar to others. Sectors are similarly represented (see appendix table 9). Exiting firms are localized in districts with the highest number of firms. On average, they are a bit smaller with 15 workers, but the median (6 workers) is equivalent.

<sup>26</sup>see the percentage of entries and exits per year in appendix table 2.D.1.

<sup>27</sup>We performed robustness checks for all estimations in the chapter dropping the largest 1% of firms.

<sup>28</sup>In data, the president is the firm's CEO. Most of time he is also the firm's owner.

<sup>29</sup>87 firm president origins are observed : Afghanistan, Albania, Argentina, Australia, Austria, Azerbaijan, Bangladesh, Belgium, Belize, Bolivia, Bosnia and Herzegovina, Brazil, Bulgaria, Canada, China, Colombia, Croatia, Czech Republic, Denmark, Egypt, Finland, France, Georgia, Germany, Ghana, Greece, Hungary, India, Iran, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, Kosovo, Kuwait, Kyrgyzstan, Latvia, Lebanon, Lithuania, Madagascar, Malaysia, Maldives, Nepal, Netherlands, New Zealand, Nigeria, Norway, Oman, Pakistan, Palestine, Peru, Philippines, Poland, North Macedonia, Russia, Saudi Arabia, Serbia, Sierra Leone, Singapore, Slovakia, Slovenia, South Africa, South Korea, Spain, Sri Lanka, Swaziland, Sweden, Switzerland, Syria, Tajikistan, Thailand, Turkey, Turkmenistan, Ukraine, United Arab Emirates, United Kingdom, United States of America, Uruguay, Uzbekistan, Yemen, Zimbabwe.

<sup>30</sup>Including agriculture, formal labour represents less than 10% of total employment in Afghanistan (Bonnet et al.,

the biggest, most capital-intensive and productive firms and are the state's main revenue source (Bonnet et al., 2019; Mashal, 2014). Furthermore, we can observe a strong expansion of the formal economy over the period, whose number of firms has been almost multiplied by 10 since 2003 (see appendix graph 2.D.1).

On average firms are exposed to 82 military events over the whole period. Yet, 25% of them are exposed to less than 6 events, and 50% to less than 59. In figure 2.2.1.b we represent the number of military events by district. Military events are also well dispersed throughout the country, even though their location seems mainly driven by population, geography and ethnic groups. Indeed, events are more concentrated in most populated areas, with a high but not extreme altitude (see appendix figure 2.D.2 for details on altitude and major cities in Afghanistan). Furthermore, we observe fewer events in the northern part of the country, where clans are historically more opposed to the Taliban.

Comparing figures 2.2.1.a and 2.2.1.b, the spatial correlation between the military events and the number of formal firms in districts is not clear. However, on graph 2.2.2.a we compare the empirical distribution of the number of formal firms by district depending on military intervention intensity (top vs bottom quartiles). Additionally, we display on graph 2.2.2.b the empirical distribution of firms' lifespan depending on whether they have been strongly exposed to military events (top quartile) or poorly (bottom quartile). On each graph, we can notice a clear positive correlation with exposure to military events.<sup>31</sup> In a logic of state-building process with competition for territory control, this correlation suggests that, in the Afghan case, military events capture the investment from the state to bring security and control over the area. Yet, this may also characterize the presence rent-seeking mechanisms (Berman et al., 2017) – i.e. targeting militarily areas because of their economic activities. Therefore, these descriptive pieces of evidence already call for a careful identification strategy when using military events as a proxy for insecurity.

Table 2.2.1: Firms statistics

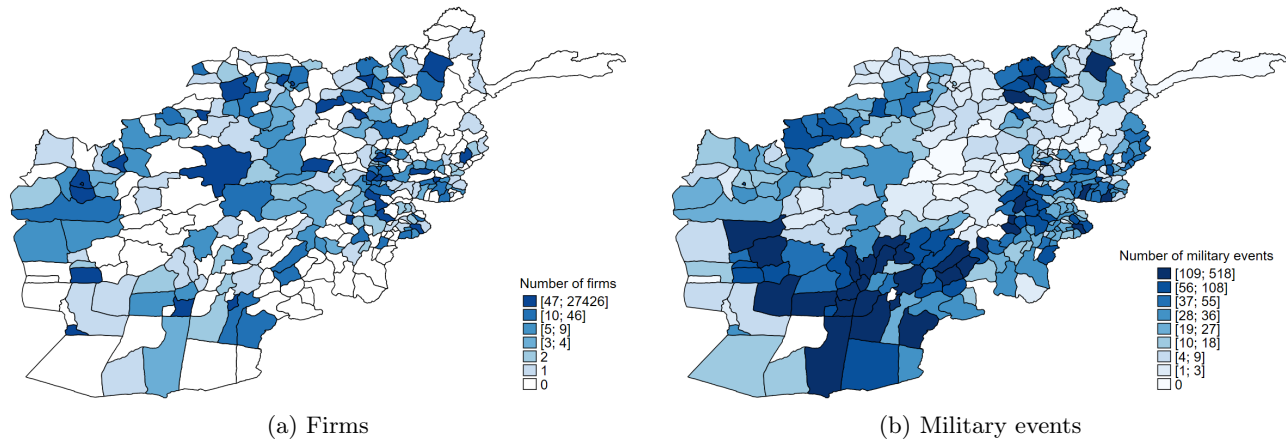
	Average	Median	Max
Size (in workers)	20	6	11928
Military events (sum)	82	59	518
Nationality (in%)	Afghan	NATO	Other foreign
	91.2	5	3.6

Authors calculation. Data are retrieved from the 2003-2016 register of formal companies of the Afghan Ministry of Economy.

2019).

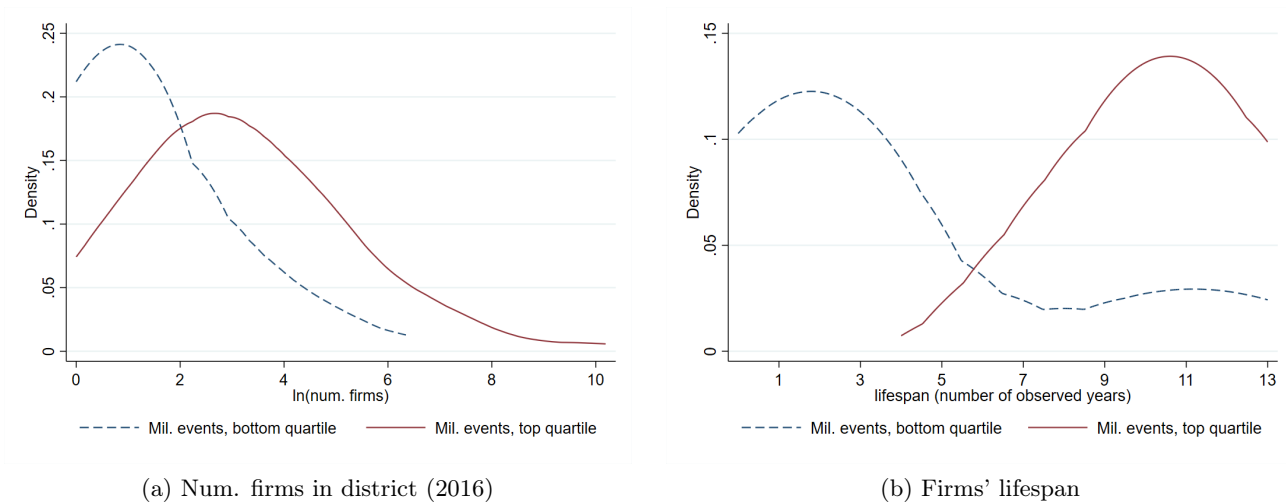
<sup>31</sup>The distribution with high military events exposure lies more on the right than the ones with low exposure.

Figure 2.2.1: Firms and events over districts (2003-2016)



Note: Classes are defined using the quantiles method.

Figure 2.2.2: Firms and military events correlation



Note: K-density graphs (a) is realized using district level number of firms and military events over the period. K-density graphs (b) displays firms' number of formally active years over our period (max 13) depending on their district's exposure to military events.

## 2.3 Exposure to military events

### 2.3.1 Identification strategy

This section presents the empirical strategy adopted to test the effect of exposure to military events on the formal activity probability and discusses its quality as a proxy for insecurity. Based on equation (2.8), we estimate  $\varepsilon_{\Gamma}(Mil.ev.it) = \varepsilon_{\Gamma}(\psi_{it}) * \varepsilon_{\psi}(Mil.ev.it)$  – the effect of the exposure to military events on the formal activity probability through changes in insecurity cost. In the Afghan context, we expect military events to mainly capture a state-capacity signal (Berdal, 2017; Fritz and Menocal, 2007; Rocha Menocal, 2011; Thies, 2007), and therefore the average change in insecurity and formal activity probability to be positive.<sup>32</sup> Yet, these events may be highly heterogeneous like their effect, which we discuss a second time. Since the formal activity probability expression (eq. 2.8) is non-linear, we perform a logit estimation. Our baseline specification is as follows:

$$\ln\left(\frac{\Gamma_{it}}{1 - \Gamma_{it}}\right) = \beta Mil.ev.i,t-1 + \lambda_i + \lambda_t + \epsilon_{it} \quad (2.9)$$

where  $\lambda_i$  the firm fixed effects, and  $\lambda_t$  the year ones. The formal existence probability  $\Gamma_{it}$  is approximated with a dummy variable taking the value one if the firm exists formally at time  $t$  and zero otherwise.

$Mil.ev.i,t-1$  measures the intensity of the exposure to military events. It captures whether the firm is treated (a military event has happened in the firm’s district), and the magnitude of the treatment (the number of events). Once assigned to the treatment group, the firm keeps this assignment for the next periods. Yet, intensity can increase over time if new events occur. In this difference in difference setting, we therefore estimate the shift in trend induced by a variation in the exposure to military events.

In the literature, the effect of conflict events is commonly estimated considering observations as treated only at the year of occurrence.<sup>33</sup> In our identification strategy, we choose instead to keep firms assigned to the treatments groups for two main reasons. First, we do not want to measure only the short-term effect of military events, but the general one. In the Afghan context, military events may contribute to the state-building process by expanding state control over districts. This potential source of insecurity reduction for formal firms may take time to be effective. Therefore, focusing only on the year of occurrence would omit this potential long term insecurity variation.

<sup>32</sup>In this chapter, we focus on the direct effect of insecurity. As shown by Couttenier et al. (2022), in case of firms connexions through inputs-outputs, the effect of insecurity changes can be amplified due to indirect exposure. In the Afghan economy we do not expect huge connexions, nonetheless, such indirect exposure may cause identification issues in the presence of selection bias. We address this in a further section using an Inverse Propensity Weighting (IPW) strategy.

<sup>33</sup>For instance: Camacho and Rodriguez (2013); Couttenier et al. (2022); Del Prete et al. (2023); Klapper et al. (2013).

Second, considering observations as treated only at time  $t$  imposes we assume that the treatment has only an effect in  $t$ . If this assumption does not hold, which is very likely in our case, and treatment is not staggered, results will be biased. This issue has been deeply treated by the literature on twoway-fixed-effects estimators (De Chaisemartin and d’Haultfoeuille, 2020). Limiting treatment to time  $t$ , the estimator uses observations treated in the past as control and compares the treatment group with a control group including observations whose trend has already shifted due to previous treatment. Our design avoids this issue.

Military events exposure is not random. We already mentioned that, because of rent-seeking behaviour, we may face reverse causality concerns. To address this  $Mil.ev_{i,t-1}$  is lagged – only events anterior to the observation in  $t$  are considered. Additionally, our fixed effects capturing any year, district or firm-level variable, they address a large sample of potential omitted variables, such as geography, ethnicities, infrastructures, firms’ nationality or total expenditures. The only remaining possible source of omitted variable bias is time-varying district or firm-level shocks that may affect both  $Mil.ev_{i,t-1}$  and  $\Gamma$ . are either country-time, district or firm-specific and, therefore, they should be captured by the fixed effects. Yet, empirically, we may imperfectly control for some variables, especially the wage, which, because of unachieved free movement, may not be only different over time, but also over districts. Consequently, we further address endogeneity concerns with an Inverse Propensity Weighting (IPW) in section 2.3.3, and an Instrumental Variable (IV) approach in section 2.4.

### 2.3.2 Baseline Results

Results are displayed in table 2.3.1. In the first column, the coefficient is the estimate of the elasticity  $\varepsilon_{\Gamma}(Mil.ev_{it})$  (Kitazawa, 2012; Silva and Kemp, 2016). In the second one, we report odd-ratios.<sup>34</sup> In both estimations, we observe a significant positive effect of exposure to military events on the formal activity probability. Interpreting column 1 results, we find that an increase of 1% in the exposure to military events raises the probability of formal existence by 0.211%.

Yet, military events may have dynamic and heterogeneous effect over time and districts. In this case, the Average Treatment on the Treated (ATE) may suffer from negative-weight bias (De Chaisemartin and d’Haultfoeuille, 2020). To address this issue, we perform a robustness check using De Chaisemartin and d’Haultfoeuille (2024) estimator. The results confirm the absence of such bias (see table 2.3.2).<sup>35</sup>

<sup>34</sup>The logit-fixed-effects estimator allows for robust standard errors an for clustering but only at the same level as the group variable fixed effect – here the firm level. Yet, we may also expect spatial autocorrelation – several firms facing the same district level exposure to military events. Therefore we replicate our estimation with district-level clusters and results are reported in odd-ratios. Furthermore, Kwak et al. (2023) showed that in the presence of serial autocorrelation the logit estimator is inconsistent even using individual/group clusters (in our case firm clusters). We address address this issue by using the De Chaisemartin and d’Haultfoeuille (2024) robust-did linear estimator with firm clusters (results are reported in section 2.3.4 All specifications lead to the same conclusion.

<sup>35</sup>We further discuss the dynamic and heterogeneous effect of military events in section 2.3.4.

Thus, on average, military engagement in Afghanistan positively impacted the formal economy. This supports the prevalence of a state capacity signal, meaning that through its military interventions, the state expends its control over districts and brings more security to formal firms. Yet, compared to the size we should expect from the theoretical elasticity in case of a decrease in insecurity costs ( $\varepsilon(\psi) = -2\theta - k(\sigma - 1) \in [-11.86; -7.86]$ ), the estimated coefficient is very weak. This suggests that an increase in exposure to military events does not capture clear changes in insecurity (see section 2.1.4).

Table 2.3.1: Military events exposure

Estimator: Dependent variable:	Logit	
	Firm's formal existence elasticities (1)	odds ratios (2)
Military events	0.211 <sup>a</sup> (0.064)	1.527 <sup>b</sup> (0.309)
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	479.150	479.150

FE, Fixed effects; standard errors are in parentheses; columns (1) and (2) standard errors are respectively clustered at the firm and district levels. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively

### 2.3.3 Inverse Propensity Weighting

The exposure to military events being not random, our results may suffer from selection bias: changes in exposure may be endogenously assigned to firms, leading to incomparable treatment and control groups. Therefore, we test the sensitivity of our baseline to Inverse Propensity Weighting (IPW) (Austin and Stuart, 2015; Lunceford and Davidian, 2004). This approach relies mainly on the same assumptions as propensity score matching – the most important one being that by reducing the selection bias on observable variables, we also improve control for unobservables. Yet, IPWs have the advantage of not reducing the sample size.

We define the propensity score  $p_i(x)$  as the likelihood of experiencing a change in the exposure to military events between 2003 and 2016 conditionally on a set of firm characteristics.<sup>36</sup> We compute two propensity scores. The first one, which we call "standard", is conditional on the firm's size (in workers) and (president) nationality. The second score, which we call "extended", adds firm sectors.<sup>37</sup> From each score, we determine firms' *inverse probability weights* defined by  $w = \frac{T}{p_i(x)} + \frac{1-T}{1-p_i(x)}$  (Austin

<sup>36</sup>Therefore, firms that have a positive exposure in 2003 which is stable over the period are in the control group.

<sup>37</sup>Sectors *Activities of extraterritorial organisation, Activities of extraterritorial organisation, Public administration and defence; compulsory social security and Real estate activities* are dropped because of collinearity issues.

and Stuart, 2015; Lunceford and Davidian, 2004).<sup>38</sup> We display IPW estimations results in table 2.D.3 column 1 and 2. Compared to the baseline,  $Mil.ev_{i,t-1}$  coefficients are very similar even if slightly superior. In the end, the IPW strategy supports the robustness of our baseline results.

### 2.3.4 Treatment heterogeneity

The small size of the elasticity estimated in the baseline may result from composition issues in  $Mil.ev_{i,t-1}$  – namely the potential aggregation of events with opposite effects on formal firms’ insecurity costs. In this section, we hence analyse the heterogeneity behind the average effect of the exposure to military events on firms’ formal existence. We discuss five potential sources of heterogeneity: time-related heterogeneity, cooperative vs external events, events’ intensity, events’ concentration and provincial events.

*Time-related heterogeneity*– In our context, military events have, on average, a positive effect of formal activity. Yet, their effect may be dynamic. Especially, at the year of occurrence, military events may have a negative effect because of rise in conflict intensity, while over time may only subsist a positive effect due to an expansion of the state capacity. To test this potential difference between short-term and long-term effects, we include in our baseline estimation a new variable,  $Mil.ev.shocks_{i,t}$ , summing all events the firm is exposed (only) at time  $t$ . While  $Mil.ev_{i,t-1}$  captures time lasting change in insecurity,  $Mil.ev.shocks_{i,t}$  captures short-term insecurity variation. Results are reported in appendix table 2.D.3 column 3. They confirm the existence of an average negative effect at the year of occurrence which then becomes positive.

Additionally, we include an interaction variable between  $Mil.ev.shocks_{i,t}$  and  $Mil.ev_{i,t-1}$ . Its coefficient captures the dependence of the military events’ short-term effect (at time  $t$ ) on the past exposure (from 2003 to  $t-1$ ). Results, in column 4, show that military events shocks have a substantially more negative effect in areas with low past exposure (i.e. that were mostly spared by conflict).

Performing an event study, we directly estimate the dynamic effect of military event. Dynamic effect estimations are particularly sensitive to negative-weight bias. Using De Chaisemartin and d’Haultfoeuille (2024) estimator, we address this issue. We observe in table 2.3.2 that, even after the first period, the military events’ effect strongly grows over time. Therefore, our results support that the reduction in insecurity is very progressive.

*Cooperative vs external events*– Since foreign states are not politically or financially dependent on

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<sup>38</sup> $T$  is a dummy taking value one if the firm experience a change in the exposure to military events. We test whether under observations’ weights treated and control groups are comparable. We test whether under observations’ weights treated and control groups are comparable. We follow Rubin (2001); Stuart and Rubin (2008) and consider the standardised difference in means (B) and the variance ratio (R). To conclude that the groups are comparable, B must be inferior to 0.25, and R between 0.5 and 2. With the standard IPW we obtain  $B=4.8<15\%$  and  $0.5<R=1.18<2$  while with the sector extended one  $B=13.6<15\%$  and  $0.5<R=0.9<2$ . Thus, in both sets of matched observations, the control and treated groups are well balanced.

Table 2.3.2: Chaisemartin D'Haultfoeuille (2024) estimator

Dependent variable:	Firm's formal existence					
Time	t	t+1	t+2	t+3	t+4	t+5
Military events	0.036 <sup>a</sup>	0.081 <sup>a</sup>	0.130 <sup>a</sup>	0.210 <sup>a</sup>	0.300 <sup>a</sup>	0.400 <sup>a</sup>
Standard error	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
No. observ.	88.011	83.979	52.133	51.041	50.232	49.351
No. switchers	31.967	31.955	31.954	31.954	31.954	31.948

*Note:* t is the year of the change in events. For each period the reference is t-1. Because our treatment is continuous, the estimator requires we set an "insignificant level of treatment" – observation with an inferior treatment intensity are used as control. We set this threshold to 2 events. Bootstrapped standard errors clustered at the firm level are in parentheses. a, b and c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

local formal firms, they may be less incited to improve their security. Using the detailed information on military events in our data, we divide these events into two categories: the cooperative events, which imply a co-belligerence between the foreign and the local states (e.g. ISAF & Afghan state vs the Taliban), and the external events, in which only foreign countries are involved against a non-state organised actor (e.g. US vs the Taliban).<sup>39</sup> Then, we replicate the baseline estimation but distinguish between the exposure to cooperative events and to external ones. Results are presented in appendix table 2.D.3 column 5 and 6. The cooperative events' coefficient is very close to the baseline one, while the external events' coefficient is negative and statistically not significant. Thus, this supports that our baseline results are driven by cooperative events.

*Military events' intensity* – Military events do not all have the same intensity. We can approximate this intensity using the number of deaths associated with each event. Therefore, we replicate our baseline estimation but substitute deaths for military events. We construct the variable  $Deaths_{it}$  as  $Mil.ev.it$ . Yet, by summing events' deaths,  $Deaths_{it}$  weights military events depending on their number of deaths.<sup>40</sup> The number of fatalities is expected to capture the level of destruction associated with events, but also the magnitude of the events in general. Hence, an increase in  $Deaths_{it}$  may also express a stronger state capacity signal. We present the estimation results in the appendix table 2.D.4 column 1. The coefficient of  $Deaths_{it}$  is very close to our baseline.<sup>41</sup> This supports that once accounting for events' intensity, the state-capacity signal effect still prevails over the destruction effect. To better disentangle these two dimensions in our measure of military events' intensity, we run a second estimation dissociating *soldier* deaths from *civilian* deaths. Soldier deaths should better capture the state capacity signal. On the contrary, civilian deaths should better proxy events' damages. Results in table 2.D.4 column 2 confirm this assumption. Soldier deaths' coefficient is positive and superior to

<sup>39</sup>In our data, no event involves only the local state. Therefore, all observed military events are either cooperative or external. In the majority, the military events are cooperative. Nonetheless, external events are not negligible since 7% of firms have been exposed to them.

<sup>40</sup>In our data, military events have, by definition, at least one death. Thus, this proxy gives to all events a positive intensity. Observations with observed military events but no declared death are dropped.

<sup>41</sup>Because  $Deaths_{it}$  and  $Mil.ev.it$  are too collinear we cannot include them both in the estimation. Furthermore,  $Mil.ev.it$ 's coefficient would not be interpretable since it would correspond to a fake event with no intensity.



$Deaths_{it}$ 's one, while civilian deaths' coefficient is negative.

*Events concentration*– Agents do not only react to direct exposure to events, but also to the risk of exposure and its uncertainty (Tapsoba, 2023). Therefore, events that deviate from the usual location and time of occurrence may be unexpected by agents and have more negative effects. To test this second assumption, we replicate our baseline estimation but weight events depending on their distance to the district-level average location (table 2.D.4 column 3) and depending on their distance to the district-level average date of occurrence (table 2.D.4 column 4).<sup>42</sup> Although, in the baseline, exposure to military events has a positive effect, exposure to the weighted events has a significant negative effect. This confirms that unpredictable events have negative consequences on formal firms and do not contribute to the reduction in insecurity.

*Provincial events*– Military events localised in a district may be correlated with events in the same provincial area reducing the insecurity cost of firm  $i$ 's competitors.<sup>43</sup> Following our theoretical model, this would imply a reduction in competitors' price and the entry of new competing firms. For firm  $i$ , this means a decrease in the price index (P), market share, profit and, mechanically, in formal activity probability (see eq. 2.8).<sup>44</sup> To address this potential indirect effect of military events, we create the variable  $Provincial.mil.ev_{i,t-1}$  summing all the military events which occurred in firm  $i$ 's province from year 2003 to  $t - 1$ , excluding events in firm  $i$ 's district. We then replicate the baseline estimation including this new variable. We display results in appendix table 2.D.4 columns 5 and 6.  $Mil.ev_{i,t-1}$ 's is slightly superior to the baseline, while  $Provincial.mil.ev_{i,t-1}$ 's coefficient is significantly negative. This contends that military events have a positive direct effect on firms' formal existence associated with a negative indirect effect through an increase in competition.

The five investigated sources of treatment heterogeneity have shown the existence of military events with opposite effects. Yet, the magnitude of the estimated coefficients is still small and none permit to conclude that one of these mechanisms drives a composition issue sufficiently important to explain the size of our baseline results. In next section, we finally tackle this issue and identify insecurity's effect on formal firms' existence.

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<sup>42</sup>Formally, the distance-to-location weight is the difference (in absolute value) between the district-level average location of events (in longitude and latitude) and the event's location (in longitude and latitude) plus one. The distance-to-date weight is the difference in number of days between the yearly district-level average date of events' occurrence and the date of the event. We tested alternative weights using the same information. Results are poorly affected.

<sup>43</sup>Administratively districts are included in provinces.

<sup>44</sup>Following our theory, the price index is common to all firms. Yet, in a developing country like Afghanistan, where markets may not be fully integrated, we may observe firm or district/province specific price indexes variations that the time-fixed effect cannot absorb.

## 2.4 The reduction of insecurity: identification and impact

### 2.4.1 Instrumental variable strategy

To draw conclusions on insecurity’s impact with external validity, two remaining issues must be addressed. First, the actor initiating a military event may strongly matter for formal firms. As discussed previously, in our context of state-building, the Afghan state is a security provider for formal firms, while, on the contrary, the Taliban represent a threat (Barfield, 2010; Maley, 2020; Rocha Menocal, 2011; Thies, 2007) Yet, in events data, information on initiating actors is not available.

Second, there may persist endogeneity. We already discussed and addressed several potential address, or residual time-varying district- or firm-level omitted variables that fixed-effects do not capture (such as aids or subsidy shocks). To deal with these issues, we develop an instrumental variable (IV) strategy. In this way, we are able to select the events variations from a unique belligerent and can use the presence of external actors in the Afghan context to estimate an exogenous change in insecurity.

To build our IV, we follow Bartik (1991)’s *shift-share* approach. We interact two variables: the sum of NATO military events outside Afghanistan<sup>45</sup> (shift) and an altitude<sup>46</sup> ratio comparing the district level to the average altitude in Afghanistan (share). Thus, the instrumental variable takes the form:

$$IV_{dt} : \ln(Mil.ev.out_t) * \frac{\ln(altitude_d)}{\ln(altitude)} \quad (2.10)$$

The first variable is time-varying. Exploiting the presence of foreign actors, it exogenously defines the yearly stock of events in Afghanistan using a measure of the global interventionism of NATO in the rest of the world – the more NATO has an interventionist stance, the more it should be willing to intervene in Afghanistan. The second variable is district-specific and determines the allocation of events across space. More events should be allocated to districts whose average altitude is higher. High-altitude lands offer higher military advantages (control of accesses and valleys, strategic frontiers, protective and performable landscapes, better information, or suitable operations bases). This is particularly true in Afghanistan where mountainous areas are key to control connections between cities and between frontiers (see appendix figure 2.D.2). Furthermore, *ISAF Placemats Archives*<sup>47</sup> show that troops are mainly concentrated in the Est region and in the northern part of the Central region, where the altitude is the highest, investing less the western and northern parts of the country where the plains are more abundant (Fetzer et al., 2021).

<sup>45</sup>As for  $Mil.ev.i,t$  we do no assumption on events dynamic effects. Therefore the variable sums all events from year 2003 to  $t$ . See section 2.3.1 for a proper discussion on using continuous treatment.

<sup>46</sup>Geographic data are retrieved from the World Bank’s Afghanistan Spatial Database.

<sup>47</sup>ISAF Placemats Archives are available from 2007 to 2014.

The IV has to respect the exclusion restriction to be valid. In other words, a change in  $IV_{dt}$  must affect the firm's formal existence only through the exposure to military events. By definition, geography is not determined by human activities, while outside events are exogenous to the Afghan context. Furthermore, firms and time fixed effects ensure that independent effects of altitude or outside events are captured. Hence, we are confident of the IV's ability to capture only military event variations driven by factors exogenous to the local context and economic considerations. Still, in section 2.4.4, we further discuss the IV validity. Using alternative IV estimations<sup>48</sup> and the plausible exogeneity test proposed by Conley et al. (2012), we confirm the robustness of our approach.

Again, our model being non-linear, we use the log of the sum of outside events. Because variations in altitude across extremely elevated districts weakly affect the allocation of events (see appendix table 2.D.5), we also use the log of altitude, giving more weight to variations in altitude across less elevated districts. We perform a robustness check using untransformed altitude and dropping districts with very high altitude (the top quartile). Both approaches lead to the same results (see section 2.4.4 for more details).

The IV relies strongly on the existence of effective foreign military involvement in Afghanistan. Nonetheless, based on NATO declarations,<sup>49</sup> ISAF as followed a disengagement process since 2009. In its final phase, in 2013, this process aimed to switch from direct military involvement of NATO to a more distant support (Fetzer et al., 2021). Testing this disengagement, it appears that, indeed, after 2012 the capacity of the IV – based on NATO interventionism – to explain military events is strongly weakened (see appendix table 2.D.6). Because of this context, we can mobilise this exogenous source of events variation only before 2013. Therefore, we perform our instrumental variable estimations using a panel stopping in 2012.

Besides addressing endogeneity concerns, our IV strategy constrains the military events' to vary depending a unique belligerent: NATO. Knowing that NATO is allied with the Afghan state and almost never intervenes militarily without its involvement,<sup>50</sup> these military events variations should expand state capacity and reduce formal firms' insecurity. By using the IV, we focus on events caring a state capacity signal and exclude variations driven by the Taliban which should have an opposite effect. In this way, we overcome the lack of information on initiating actors in data and improve the identification of a clear change in insecurity.

We use an OLS/logit two-stage estimator. With an OLS, we estimate the effect of the IV on  $Mil.ev.it$ . Then, depending on this causal relation, we instrument  $Mil.ev.i,t$  and estimate its effect on

<sup>48</sup>We test the IV validity controlling for provincial events, altitude in level, with lagged variables, excluding year 2003 as well as Kabul district and using ruggedness instead of altitude.

<sup>49</sup>Source: Nato publication *Integal: Transition to Afghan Lead*.

<sup>50</sup>Some members of NATO can intervene alone. Yet, these events are rare and do not drive our results. See discussion on cooperative events in section 2.3.4.

the formal activity probability with a logit. In both stages, we include firm and year fixed-effects.

## 2.4.2 Results

We display the results in table 2.4.1. Quantifying the theoretical expression of the insecurity elasticity (see section 2.1.4), the identification of a clear insecurity change through military events should lead to a coefficient significantly larger than in our baseline estimation. Furthermore, since our IV strategy focuses on state military events by restricting the source of variation in events to states, we should observe a reduction in insecurity and therefore a positive effect on firms' formal existence. Results follow our expectations. The instrument variable has a statistically significant positive impact on  $Mil.ev.it$ . When the number of NATO events outside Afghanistan rises by 1%, the number of military events rises by 1.074%. Additionally, LM and F-stats exclude weak-IV concerns. In the second stage, we observe a sizeable positive coefficient. When the firm's exposure to instrumented military events rises by 1%, its formal activity probability rises by 4.16%. Translated in insecurity cost, this result is interpretable as the causal effect of an insecurity cost reduction between 0.35% and 0.53%.<sup>51</sup>

When the second-stage estimator is non-linear, Lin and Wooldridge (2019) recommend a control function approach for consistency reasons. Instead of using the predicted number of events, we include first-stage residuals as control. Results, reported in appendix table 2.D.7, confirm the OLS/Logit twostage conclusions. Furthermore, the coefficient of the first-stage residuals shows us that our IV strategy properly excludes negative estimated effects provoked by non-state military events or omitted variable bias.<sup>52</sup>

## 2.4.3 Firms heterogeneity

In the previous section, we estimated the causal effect of insecurity through state military events. The observed treatment is at the district level. Therefore, we assumed that all firms face the same exposure to military events inside the district. We now relax this assumption. Indeed, some micro-heterogeneity in the treatment allocation may exist. Especially, states may protect firms differently depending on their characteristics (Berman et al., 2017). First, we study the dependence on the firm's productivity. How the firm's productivity may change the treatment allocation is not evident. The state may focus less on productive firms, since insecurity costs threaten less their existence. But, on the contrary, it may protect them more since they contribute more to its finances. We do not directly observe firms'

<sup>51</sup>Considering the state military events elasticity  $\varepsilon_{\Gamma}(Mil.ev.it) = 4.16$  from table 2.4.1, and making the ratios between this elasticity and the insecurity cost elasticity thresholds  $-7.86$  and  $-11.86$  (see section 2.1.4), we derive that an increase in exposure to state military events by 1% decrease on average firms' insecurity costs between 0.53% and 0.35%.

<sup>52</sup>In the second stage, the logit estimator drops observations whose outcome (here formal existence) is constant. We replicated the IV estimation dropping these observations also in the OLS first stage. Results are poorly affected. Yet, in our standard IV approach we keep them since they slightly improve the IV precision.

Table 2.4.1: IV estimation

Dependent variable:	Firm's formal existence
Second stage	(1)
Estimator:	Logit
Military events	4.16 <sup>a</sup> (0.920)
Year FE	yes
Firm FE	yes
No. observ.	233.590
First-stage	
Estimator: OLS	
Instrumented variable:	Mil. ev.
IV	1.025 <sup>a</sup> (0.240)
Year FE	yes
Firm FE	yes
No. observ.	373.080
KPW rk F-stat:	18
KPW rk LM-stat:	8
AR Wald test:	10

OLS Ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; first stage standard errors are clustered at the district-year level; second stage standard errors are clustered at the firm level. In second stage, the logit-FE estimator excludes observations with a constant outcome – in our case firms, that always or never exist from 2003 to 2012. Clustering the second-stage standard errors at the district level does not affect results. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

productivity. Therefore we use their size in workers (observed only at the creation) as a proxy.<sup>53</sup>

Second, we analyse the dependence on the firm’s nationality. In the Afghan context, state military events are always performed with foreign states. By using this position at their advantage, these states could be willing to protect the firms with their nationality as a priority and allocate military events inside districts in function. We measure the firm’s nationality through the nationality of its president. Nationalities’ repartition is described in section 2.2. We create the dummy variable *NATO*, which takes the value one if the firm’s president has the nationality of a NATO member and 0 otherwise.

We replicate our IV approach but interact military events, as well as the IV, with the firm’s size (in log) and then with *NATO*. We display the results in table 2.5.1. In column (1), we only allow for productivity heterogeneity. In column (2) productivity and nationality heterogeneities are estimated. The relevant instrumental variable is positive and statistically significant in every first stages. The estimated independent effect is strongly positive. On the contrary the effect depending on productivity (size) is significantly negative: the more productive (the bigger), the lower the insecurity costs reduction. Nonetheless, with such coefficients no firm, even the biggest, faces an increase in insecurity because of its exposure to state military events. In addition, we observe that firms with the nationality of a NATO member receive a higher decrease insecurity cost. These results confirm the presence of micro heterogeneity depending on the firm’s characteristics. They support a focus on more vulnerable firms and favouritism from NATO towards firms of its members. Therefore, in Afghanistan, district exposure to state military events led to an expansion of the formal economy with a larger share of unproductive firms and foreign firms. This shows that security policies enforcing heterogeneous insecurity reduction can strongly affect the supply structure and its long term efficiency.

#### 2.4.4 IV validity

##### Alternative IV estimations

The instrumental variable’s *shift-share* construction highly relying on the Afghan context, the sensitivity of our strategy to alternative specifications needs to be carefully discussed. First, we redefine the Instrumental Variable using no longer the altitude in log but in value ( $IV = \ln(Mil.ev.out) * \frac{altitude}{altitude}$ ).

To address that the variations in altitude across extremely elevated districts poorly determines the al-

<sup>53</sup>There may be a correlation between the proxy precision and the size. Indeed, our data on size in workers are the workers declared to officials by formal firms. So, we observe only formal workers. As we developed in theoretical section 2.B.3, in presence of informal and formal labour, low productive firms mix these two kinds of labour, leading to under reported sizes for small firms. Additionally, running a QQ estimation à la Gabaix and Ibragimov (2011), we observe a concave relation between the rank and the size in log, suggesting that, effectively, the declared size of small firms is lower than the true size matching with the linear curve we should observe (more details on the QQ estimation are provided in section 2.C). Thus, knowing we have a negative correlation between the observed size and the precision of the data, we chose to drop firms whose size is too low to be unbiased. Yet, a large part of our sample are small firms. So, we must make trade-off between the quality of the size variable and the number of observations. The strongest concavity of relation between  $\ln(rank - 0.5)$  and  $-\ln(size)$  concerns firms whose size in log is inferior to 2. Therefore, by doping firms with a size inferior to 6 ( $-\ln(6) = 1.8$ ), the median, we should deal with the bias in data on size while keeping a sample sufficiently large and diverse.

location of military events (see appendix table 2.D.5), we drop those districts with a very high altitude (the top quartile). This alternative IV improves the statistical power in the first stage but does not affect second stage results (see table 2.5.2 column 1).

Second, as in section 2.3.4, we control for the provincial events. Thereby, we ensure we only capture the direct effect of an insecurity reduction. We report results in table 2.5.2 column 2. The power and precision of the IV is largely improved while military events' coefficient remains highly positive and significant. The reduction in the military events' coefficient is simply due to negative correlation between the IV and the provincial events.

Third, we add a one-year lag to the IV and the military events exposure. In this way, we further address any residual reverse causality concerns. Results are displayed in table 2.5.2 column 3. They are not very sensitive to these new specifications. The only change is a slight rise in the second-stage coefficient, which does not affect our conclusions.<sup>54</sup>

Fourth, we test the sensitivity of our IV strategy substituting outside NATO troops for outside military events. To construct the measure of the number of NATO countries' soldiers who are not deployed in Afghanistan, we use the military personnel data from the Correlate of War project (Singer et al., 1972; Singer, 1988), to which we subtract the number of troops in Afghanistan declared in the ISAF Placemats Archives.<sup>55</sup> By involving their troops in other territories, NATO should reduce its presence in Afghanistan. Therefore, this alternative IV should have a negative effect on the number of military events. We present the results in table 2.5.2 column 4. As expected, the first stage coefficient is negative and the instrumented military events' coefficient is significantly positive, which supports the robustness of our instrumental variable approach. Yet, ISAF Placemats Archives's information about NATO troops does not aim to give the exact number of soldiers but to represent the cost of the war. Therefore, this alternative IV is not completely independent from the local dynamic of the conflict and does not perfectly isolate events carrying a state-building signal, which induces a downward bias in the second stage coefficient.

Fifth, we restrict our panel by dropping the year 2003 and Kabul district while we adapt our IV to this new sample. In the previous estimations, we considered that in 2003 an identifiable Afghan state is involved in military and security actions. This point is important since the identification of the military events relies on the existence of this state. Yet, the Afghan constitution was enacted only in 2004. Therefore, we test the sensitivity of our results to this state existence assumption by dropping the year 2003. Additionally, we observe in our data sample that firms located in the capital's district (Kabul) represent 73.5% of our full sample, and 65,8% after excluding very small firms (size<6). Thus, we may fear that our results capture only the capital dynamics. To address this issue, we drop all

<sup>54</sup>See section 2.3.4 for a discussion on the dynamic effect of military events.

<sup>55</sup>Using ISAF Placemats Archives' information our panel's period is restricted to 2007-2014.

observations in the Kabul district to compare changes in military events exposure only between firms in districts with more comparable economic activities. To better match this new sample, we propose another IV. It is very close to the initial one. Simply, we substitute the ruggedness (i.e. the altitude's standard deviation) for the altitude. The theoretical ideas behind those IVs are the same, but by making that little change<sup>56</sup> we better fit with the spatial allocation of states' forces of this sub-sample. The second alternative specification's results are displayed in table 2.5.2 column 5. The IV approach is still robust, and the effect of the exposure to state military events significant and positive.<sup>57</sup> In the end, these alternative specifications confirm the robustness of our results.

### Robustness to exclusion restriction relaxation

As discussed above, the validity of our instrumental variable strategy is based on the absence of a direct effect of the IV on firms' formal activity probability. Nevertheless, because of the complexity of conflict dynamics and geopolitical decisions, the IV may not perfectly respect this restriction. Hence, we test in this section the robustness of our most complete specification (cf. table 2.5.1 column 2) to a deviation from the perfect validity of the exclusion restriction (Conley et al., 2012). The degree of deviation from the exclusion restriction can be obtained by regressing  $\Gamma_{it}$  on the interest variable  $Mil.ev.it$ , its interactions with size and nationality, IVs and fixed effects. The coefficients associated with the IVs represent an approximation of the degree of deviation from the exclusion restriction (van Kippersluis and Rietveld, 2018). We obtain significant but very small coefficients for all IVs:  $\nu_{IV} = 0.117$ ,  $\nu_{IV*\ln(size)} = 0.041$  and  $\nu_{IV*NATO} = -0.029$ . We introduce these deviations in the Conley et al. (2012) Union of Confidence Interval (UCI) estimator. In this way, we obtain for each IV lower- and upper-bound coefficients. They are reported in table 2.5.3. Because this estimator is linear, we cannot interpret the bound coefficients' size but only test for estimates' signs and statistical significance under the relaxation of the exclusion restriction. Since none of the bounds crosses zero, we can safely argue that exposure to state military events has a causal effect on the formal activity probability.

## 2.5 Conclusion

This chapter extensively analyses the effect of insecurity on formal firms' existence. We develop a theoretical model describing the conditions a firm must fulfil to be active in the formal market.

<sup>56</sup>Both variation are very correlated since mountainous areas (i.e. with a high altitude) have higher variations in altitudes than plains. In other words, that change can be understood as a switch from the level variation to the within one.

<sup>57</sup>Here,  $Mil.ev.it$ 's coefficient is somewhat smaller. Yet, by reducing the time frame, we capture less variation in  $Mil.ev.it$ . Moreover, dropping firms in the Kabul district lowers the share of small firms in the sample, while  $\varepsilon_{\Gamma}(Mil.ev.it)$  is higher for small firms (cf. section 4.3). Thus, this is not surprising we observe a slightly smaller coefficient.



Introducing insecurity in this framework, we demonstrate how it affects firms' formal activity. Based on this frame, we perform an empirical analysis. Using a detailed panel of Afghan firms, we show that exposure to military events has an average positive effect on formal firms' existence. Yet, military events' effect strongly depends on their actors, timing, and location. Thanks to an instrumental variable strategy we estimate the effect of an exogenous change in insecurity. We find that an increase in the instrumented military events of 1% raises by 4.16% firms' formal activity probability. Based on our theoretical model, this result is interpretable as the causal effect of an insecurity cost reduction between 0.35% and 0.53%. Finally, using the richness of our database, we show that, depending on their size and their nationality, firms face heterogeneous treatment, which can strongly distort the supply structure.

Our findings have important scientific and policy implications. They permit a first evaluation of the military intervention in Afghanistan, showing that it ensured a small development of the formal economy but with important heterogeneities. Additionally, they stress the critical consequences of insecurity on the economy, and that careful identification is required for any analysis of insecurity claiming for external validity. This invites future research to (re)examine the microeconomic dynamic of conflicts with particular attention given to the context-dependency and the international connections of both firms and conflicts actors.

Table 2.5.1: IV estimation and firms heterogeneity

Dependent variable:	Firm's formal existence	
Second-stage	(1)	(2)
Estimator:	Logit	Logit
Military events	5.489 <sup>a</sup> (0.922)	5.457 <sup>a</sup> (0.928)
Military events x ln(size)	-0.411 <sup>a</sup> (0.048)	-0.458 <sup>a</sup> (0.048)
Military events x NATO		1.156 <sup>c</sup> (0.629)
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	115.580	115.420
First-stage		
Instrumented variable:	Mil. int.	
IV	1.063 <sup>a</sup> (0.237)	1.060 <sup>a</sup> (0.237)
SW F-stat:	37	42
SW Chi-sq	41	47
Instrumented variable: Mil. int. x ln(size)		
IV x ln(size)	1.074 <sup>a</sup> (0.137)	1.074 <sup>a</sup> (0.136)
SW F-stat:	100	105
SW Chi-sq	112	117
Instrumented variable: Mil. int. x NATO		
IV x NATO		1.052 <sup>a</sup> (0.153)
SW F-stat:		52
SW Chi-sq		58
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	186.950	186.190

OLS, ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; first stage standard errors are clustered at the district-year level; second stage standard errors are clustered at the firm level; median size is equal to 6. In first stages, only the coefficient of the interaction variable of interest is reported. Full table is available in appendices. Clustering the second-stage standard errors at the district level does not affect results. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 2.5.2: Alternative IV strategies

Dependent variable:	Firm's formal existence				
Kabul district:	yes	yes	yes	yes	no
Altitude/Ruggedness Q4:	no	yes	yes	yes	no
Years:	2003-2012	2003-2012	2003-2012	2007-2014	2004-2012
Second-stage	(1)	(2)	(3)	(4)	(5)
Estimator:	Logit	Logit	Logit	Logit	Logit
Military events	4.323 <sup>a</sup> (1.004)	2.633 <sup>a</sup> (0.632)		1.536 <sup>a</sup> (0.581)	2.38 <sup>b</sup> (1.083)
Provincial mil. ev.		-0.071 <sup>a</sup> (0.021)			
Lagged military events			5.146 <sup>a</sup> (1.059)		
Year FE	yes	yes	yes	yes	yes
Firm FE	yes	yes	yes	yes	yes
No. observ.	229,770	233,590	233,590	182,000	27,315
First-stage					
Estimator:	OLS	OLS	OLS	OLS	OLS
Instrumented variable:	Mil. ev.	Mil. ev.	Lagged Mil. ev.	Mil. ev.	Mil. ev.
$\ln(\text{Mil.ev.out}) \times \frac{\text{altitude}_d}{\text{altitude}}$	0.001 <sup>a</sup> (0.000)				
Standard IV		1.489 <sup>a</sup> (0.293)			
Provincial mil. ev.		0.087 <sup>a</sup> (0.020)			
Lagged IV			0.806 <sup>a</sup> (0.227)		
$\ln(\text{troops.out.}) \times \frac{\ln(\text{altitude})}{\ln(\text{altitude})}$				-0.265 <sup>a</sup> (0.081)	
$\ln(\text{Mil.ev.out}) \times \frac{\text{ruggedness}}{\text{ruggedness}}$					0.901 <sup>a</sup> (0.266)
Year FE	yes	yes	yes	yes	yes
Firm FE	yes	yes	yes	yes	yes
No. observ.	363,570	373,080	373,080	298,464	69,660
KPW rk F-stat:	22	26	13	17	12
KPW rk LM-stat:	12	13	6	14	6

OLS, ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; provincial military events exclude events in the district of interest; first stage standard errors are clustered at the district-year level; second stage standard errors are clustered at the firm level. *Q4* refers to the top quartile. Because of data limitations about NATO's troops deployment, we use in column 5 a panel from 2007 to 2014 to exploit the maximum of the information. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 2.5.3: Plausibly exogenous instrument

Dep. var.: $\Gamma$	<i>Union of Confidence Interval estimations</i>	
Instrumented var.	Min 95% CI	Max 95% CI
Military events	0.111	0.256
Military events x ln(size)	-0.052	-0.010
Military events x NATO	0.035	0.072

*Notes:* UCI based on  $\nu$  coefficients from a regression of  $\Gamma$  on interest variables and the IVs:  $\nu_{IV} = 0.117$ ,  $\nu_{IV*ln(size)} = 0.041$  and  $\nu_{IV*NATO} = -0.029$   
 Robust standard errors are clustered at the district-year level.

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# Appendix

## 2.A Formal existence probability composition

This appendix section details the theoretical probabilities describing firms decisions relative to formal existence and how they are affected by insecurity.

### 2.A.1 The Entry probability

A firm enters if its expected formal profit is superior to zero. Yet, to make its entry decision, it only considers the expected fixed costs  $\underline{f}$ . Therefore, the entry condition is:

$$\frac{E}{P^{1-\sigma}} \frac{(\frac{\sigma}{\sigma-1}w\alpha)^{1-\sigma}}{\sigma} - \underline{f} > 0 \quad (2.11)$$

where  $P$  is the price index,  $w$  the wage,  $E$  the total expenditures,  $\alpha = \frac{1}{\gamma}$  the marginal cost (i.e. the inverse of productivity) and  $\sigma$  the elasticity of substitution. The probability of entry can be understood as the probability that the firm draws a formal productivity that allows a profit superior to the expected fixed cost. Knowing that  $\alpha$  follows the distribution  $G = \frac{\alpha^\theta}{\bar{\alpha}^\theta}$ , we can properly write:

$$\Gamma_e(\alpha) = ((\sigma - 1)\sigma^{\frac{\sigma}{1-\sigma}})^\theta \left(\frac{E}{P^{1-\sigma}\underline{f}}\right)^{\frac{\theta}{\sigma-1}} (w\bar{\alpha})^{-\theta} \quad (2.12)$$

The entry probability is positively impacted by the real demand  $(\frac{E}{P^{1-\sigma}})$ . On the contrary, since they all raise the cost the firm must overcome to realise a positive profit, the wage ( $w$ ), the maximum marginal cost in the economy ( $\bar{\alpha}$ ) and, the minimum fixed cost ( $\underline{f}$ ) reduce the probability. Additionally, the entry probability contains also a mark-up constant determined by the CES elasticity of substitution.

### 2.A.2 (In)Formality probability

Comparing formal and informal profit functions, the firm must decide to be formal or not. The only difference between the informal and the formal profit functions is the marginal cost parameter  $\alpha = \frac{1}{\gamma}$ .

Thus, we can summarise the formality condition to:

$$\alpha < \alpha_{inf} \quad (2.13)$$

$\alpha$  following the distribution  $G(\alpha)$  we obtain the formality probability:

$$\Gamma_f(\alpha) = \bar{\alpha}^{-\theta} \alpha_{inf}^{\theta} \quad (2.14)$$

Mechanically, having a higher informal marginal cost raises the probability, while a higher marginal cost upper-bound reduces it.

### 2.A.3 Non-exit probability

Once the firm knows its productivity and has decided to enter the market, it discovers its fixed cost and must choose whether it stays or leaves the market. Because  $f^*$  follows an inverse-Pareto distribution, most firms will face a fixed cost that is very close to the expected one  $-\underline{f}$ . Thus, to stay on the market (as a formal firm), the firm must respect the condition:

$$\frac{E}{P^{1-\sigma}} \frac{(\frac{\sigma}{\sigma-1} w \alpha_i)^{1-\sigma}}{\sigma} - f^* > 0 \quad (2.15)$$

where  $\alpha_i$  is no longer an aleatory variable since the non-exit decision occurs after the firm has observed its productivity. Therefore, knowing that  $f^*$  follows  $P(F < f^*) = \frac{f^{*k}}{f^k}$ , we can derive the non-exit probability:

$$\Gamma_x(f^*) = \left(\frac{E}{P^{1-\sigma}}\right)^k (\sigma \bar{f})^{-k} \left(\frac{\sigma}{\sigma-1} w \alpha_i\right)^{-k(\sigma-1)} \quad (2.16)$$

Because we apply a profit maximisation logic again, the parameters' contribution to the non-exit probability is very close to the one in the entry probability. The real expenditure  $(\frac{E}{P^{1-\sigma}})$  positively affects the probability while other parameters, associated with production costs, reduce it. Notice that the non-exit probability does not take into account that firms will always pay a fixed cost of at least  $\underline{f}$ . Indeed, this is already captured by the entry decision. The non-exit decision can be understood as a problem of expectation where  $f^*$  is the truly required revenue to stay on the market, while  $\underline{f}$  is the value expected and engaged at the entry by firms.

### 2.A.4 Formal activity probability

Having a formal activity means that the firms (1) entered the market, (2) chose to be formal and (3) did not exit the market. Therefore, the formal activity probability is simply  $\Gamma = \Gamma_e * \Gamma_{for} * \Gamma_x$  – the multiplication of the three probabilities we described. In this way, we can express the formal activity probability as follows:

$$\Gamma = \mu \left( \frac{E}{P^{1-\sigma}} \right)^{\frac{\theta}{\sigma-1}+k} \alpha_{inf}^{\theta} w^{-\theta+k(1-\sigma)} \bar{\alpha}^{-2\theta} \alpha^{k(1-\sigma)} \underline{f}^{-\frac{\theta}{\sigma-1}} \bar{f}^{-k} \quad (2.17)$$

### 2.A.5 Insecurity-extended probabilities

Integrating insecurity costs in the model, we obtain the following insecurity-extended entry (2.18), (in)formality (2.19) and non-exit (2.20) conditions:

$$\frac{E}{P^{1-\sigma}} \frac{\left( \frac{\sigma}{\sigma-1} w \psi \alpha \right)^{1-\sigma}}{\sigma} - \underline{f} > 0 \quad (2.18)$$

$$\alpha \psi < \alpha_{inf} \quad (2.19)$$

$$\frac{E}{P^{1-\sigma}} \frac{\left( \frac{\sigma}{\sigma-1} w \psi \alpha_i \right)^{1-\sigma}}{\sigma} - f^* > 0 \quad (2.20)$$

and the probabilities:

$$\Gamma_e(\alpha) = ((\sigma-1)\sigma^{\frac{\sigma}{1-\sigma}})^{\theta} \left( \frac{E}{P^{1-\sigma} \underline{f}} \right)^{\frac{\theta}{\sigma-1}} (w \psi \bar{\alpha})^{-\theta} \quad (2.21)$$

$$\Gamma_f(\alpha) = (\psi \bar{\alpha})^{-\theta} \alpha_{inf}^{\theta} \quad (2.22)$$

$$\Gamma_x(f^*) = \left( \frac{E}{P^{1-\sigma}} \right)^k (\sigma \bar{f})^{-k} \left( \frac{\sigma}{\sigma-1} w \psi \alpha_i \right)^{-k(\sigma-1)} \quad (2.23)$$

Thus, from equations (2.20), (2.21) and (2.22) we can define the insecurity-extended formal activity probability:

$$\Gamma = \mu \left( \frac{E}{P^{1-\sigma}} \right)^{\frac{\theta}{\sigma-1}+k} \alpha_{inf}^{\theta} \psi^{-2\theta-k(\sigma-1)} w^{-\theta-k(\sigma-1)} \bar{\alpha}^{-2\theta} \alpha^{-k(\sigma-1)} \underline{f}^{-\frac{\theta}{\sigma-1}} \bar{f}^{-k} \quad (2.24)$$

## 2.B Alternative modellings of the informal sector

In this section, we present three extensions of our theoretical model. They focus on how we model informality and bring more interdependence between sectors. First, we introduce informal insecurity cost. Second, we derive the formality probability under the assumption of equivalent position in the formal and informal productivity distributions. And third, we introduce (in)formal labour.

## 2.B.1 Informal insecurity

In the standard model, insecurity costs concern only the formal sector. Now, we assume that the informal sector has its own level of insecurity and, therefore, its own insecurity cost parameter  $\psi_{inf}$ . Under this new assumption, we the formality probability is defined as follows:

$$\Gamma_f(\alpha) = \left(\frac{\psi}{\psi_{inf}}\right)^{-\theta} \bar{\alpha}^{-\theta} \alpha_{inf}^{\theta} \quad (2.25)$$

In the model, this point has no big consequences. Yet, it helps to understand the externality issues we may face estimating the formal insecurity elasticity  $\varepsilon(\psi)$ . Indeed, empirically,  $\psi$  and  $\psi_{inf}$  are potentially not completely independent and an instrument affecting one has a good chance to affect the other too. For example, if in a given territory the state deploys more police or military forces, this will increase the security of formal firms.<sup>58</sup> Yet, this may also reduce the security in the informal sector because of a higher probability of being caught and punished by the state. Thus, while estimating  $\varepsilon(\psi)$  its highly probable one may not capture only the "pure" elasticity  $(-2\theta - k(\sigma - 1))$ , but also some externalities because of plausible interdependence between the formal and informal insecurity.

## 2.B.2 Interdependent productivities

Previously we assumed that firms independently draw their formal and informal productivities. Nevertheless, this may lead to the possibility to draw productivity parameters very different between sectors. In this section, we address this issue by assuming that firms no longer draw productivity parameters but a place in the productivity distributions. In other words, we set for each firm:

$$\frac{\alpha^{\theta}}{\bar{\alpha}^{\theta}} = \frac{\alpha_{inf}^{\delta}}{\bar{\alpha}_{inf}^{\delta}} \quad (2.26)$$

Deriving the formality probability under this restriction, we obtain:

$$\Gamma_f(\alpha) = \bar{\alpha}^{-\left(\theta \frac{\theta}{1-\frac{\theta}{\delta}}\right) - \theta} \left(\frac{\psi}{\alpha_{inf}}\right)^{-\theta \frac{1}{1-\frac{\theta}{\delta}}} \quad (2.27)$$

Equation (2.25), reveals a new parameter  $-\bar{\alpha}_{inf}$  and new elasticities. Focusing on the insecurity elasticity in the formality probability, we observe that the standard elasticity  $-\theta$  (see eq. 2.21), is extended by a component capturing the interdependence between formal and informal productivities  $\left(\frac{1}{1-\frac{\theta}{\delta}}\right)$ . Knowing that we set  $\delta > \theta$ ,<sup>59</sup> this component can only increase the elasticity. Indeed, the closer the formal and informal productivities distributions' shapes, the lower the technological advantage of

<sup>58</sup>In the standard model, risk of detection and sanction of informal firms is captured by the informal productivity distribution and assumed as independent from the level insecurity in the formal sector.

<sup>59</sup>If  $\delta < \theta$ ,  $\Gamma_f(\alpha)$  would no longer capture the probability of being formal but the probability of being informal).

the formal sector, and the more matters production costs and productivity minimums. If formal and informal productivity distributions' shapes are extremely close, a slight reduction in insecurity costs will be sufficient to switch all informal firms to formality. Thus, with this simple extension, we immediately observe that the productivities' interdependence reinforced the insecurity elasticity.

### 2.B.3 (In)formal labour

Until now, we have defined informality through a sector perspective. Changes in informality that define  $\Gamma_f(\alpha)$ , and by extension  $\Gamma$ , are purely extensive – the firm switches from one sector to the other. In this section, we allow for intensive margins of informality by introducing formal and informal labour as [Ulyssea \(2018\)](#). Firms now face two kinds of labour:  $l_{inf}$  the informal labour, and  $l_f$  the formal one. Both have the same productivity directly determined by the firm's productivity as previously. Yet, they have different costs. In exchange for one unit of formal labour, firms must pay the wage  $w_f = \tau_f w$ , where  $\tau_f > 1$  is a constant payroll tax on formal workers. On the other hand, by hiring informal labour, firms avoid tax costs but face a probability of being detected and sanctioned by officials. This takes the form of a wage distortion  $w_{inf} = \tau_{inf}(l_{inf}; l_f)w$  where  $\tau_{inf}(l_{inf}; l_f) > 1$  is an expected cost assumed to be increasing and convex in firm size ( $\tau'_{inf}; \tau''_{inf} > 0$ ). These assumptions follow the stylised fact that larger firms are more likely to be detected ([De Paula and Scheinkman, 2011](#); [Fortin et al., 1997](#); [Ordenez, 2014](#); [Ulyssea, 2018](#)). Informal firms can only hire informal workers, but formal firms can hire both. Additionally, we assume that  $\frac{d\tau_{inf}}{dl_{inf}} > \frac{d\tau_{inf}}{dl_f} > 0$  – meaning that the expected cost is more sensitive to changes in informal labour than in formal one. In this way, we allow firms to be informal, formal with a mix of formal and informal labour, or formal with only formal labour<sup>60</sup>. Therefore, under these new assumptions on the labour market, we obtain the formality condition:

$$\alpha\psi\left(\frac{l_f}{l}w_fw_{inf}^{-1} + \frac{l_{inf}}{l}\right) < \alpha_{inf} \quad (2.28)$$

This extension has two interests. First, it underlines another potential source of insecurity externalities: the sensitivity of the informal labour cost to insecurity ( $\frac{d\tau_{inf}}{d\psi} < 0 \Rightarrow \frac{dw_{inf}}{d\psi} < 0$ ). Second, working in the empirical sections with official data on firms, it may be useful to keep in mind that, under such (in)formal labour structure, the lower the formal firm's productivity, the higher the probability of under-reported size.

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<sup>60</sup>If  $\frac{d\tau_{inf}}{dl_{inf}} = \frac{d\tau_{inf}}{dl_f} > 0$  there are only fully formal or informal firms. And if  $\frac{d\tau_{inf}}{dl_{inf}} > 0$  &  $\frac{d\tau_{inf}}{dl_f} = 0$  firms are either fully informal or mixing formal and informal labour, but none is hiring only formal labour.

## 2.C Quantification exercise

We now detail our estimation of  $\theta$ , the Pareto shape parameter of the formal productivity distribution. We perform a QQ regression (Head et al., 2014). Because we only observe the tail of the total productivity distribution, we follow Gabaix and Ibragimov (2011) methodology. We regress the theoretical quantiles  $\ln(\zeta - 0.5)$ , with  $\zeta$  the firm's descending order of size, on the empirical quantiles of the sorted log size  $(-\ln(\text{size}))$ .<sup>61</sup> In the CES-Pareto structure of our theoretical framework, the coefficient of such regression,  $\tilde{\theta}$ , gives us the Pareto shape parameter divided by the elasticity of substitution minus one (i.e  $\sigma - 1$ ). Thus, to recover the right estimation of the Pareto shape parameter, we multiply  $\tilde{\theta}$  by  $\sigma - 1$ , with  $\sigma = 5$  (Head et al., 2014). In the end, we obtain  $\theta = 3.93$ . The result is equivalent using Head et al. (2014) method to perform the QQ estimation. Following their methodology, we regress the empirical quantiles of the sorted log size on the theoretical quantiles (i.e.  $-\ln(1 - ((k - 0.3)/(n + 0.4)))$ , where  $k$  is the firm's ascending order of size and  $n$  the rank of the firm having the largest size. There, the coefficient gives us  $1/\tilde{\theta}$ . To recover the right estimation of the Pareto shape parameter, we must here divide  $\sigma - 1$  by the coefficient. We obtain  $\theta = 4$ .

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<sup>61</sup>In footnote 51, we underline that the quality of size as a proxy strongly decreases for very small firms. So, once again we drop those very small firms(i.e with less workers than 6 at entrance on the formal market).

## 2.D Supplementary tables and figures

Table 2.D.1: Share of firms creations and destructions by year

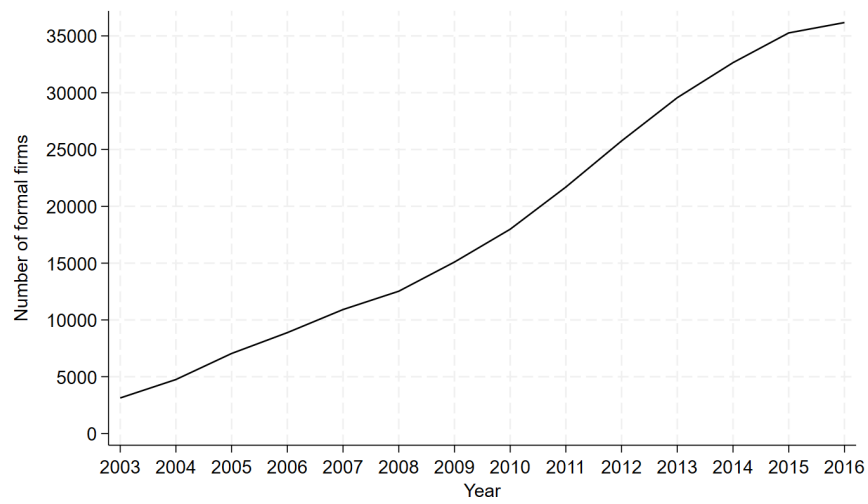
Year	Creations(%)	Destructions (%)
2003	8.38	0
2004	4.43	2.55
2005	6.23	2.64
2006	5.04	4.05
2007	5.61	5.28
2008	4.53	7.04
2009	7.07	6.16
2010	8.08	10.99
2011	10.32	11.70
2012	11.18	9.94
2013	10.53	11.08
2014	8.69	14.25
2015	7.33	10.11
2016	2.57	4.22

*Notes:* Author's calculation. In 2016, 2.57% of the total creations and 4.22% of the total destructions are observed.

Table 2.D.2: Sectors

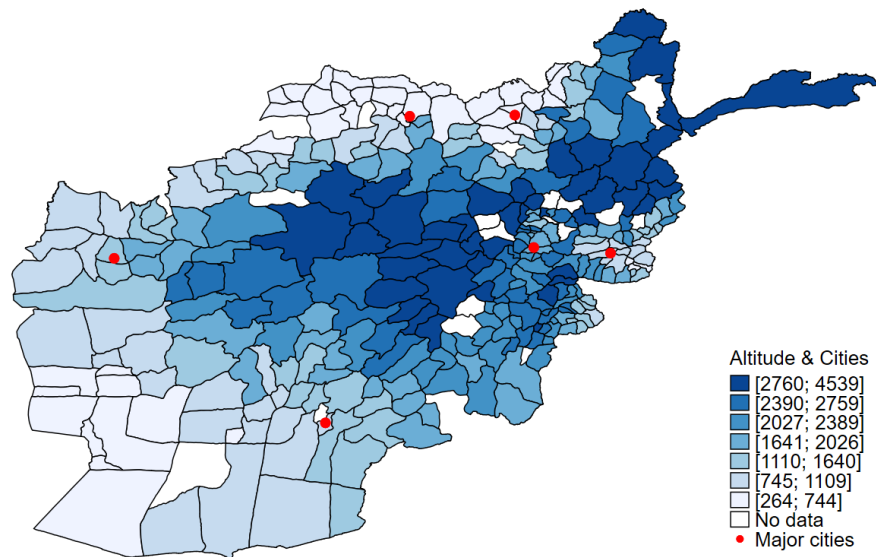
Sector	All firms		Exiting firms	
	Number	Percentage	Number	Percentage
Accommodation and food service activities	146	0.39	3	0.27
Activities of extraterritorial organisation	1	0.00	0	0
Administrative and support service activities	579	1.55	17	1.50
Agriculture, forestry and fishing	859	2.3	21	1.86
Arts, entertainment and recreation	35	0.09	0	0
Construction, and construction materials production	15,171	40.66	573	50.66
Education	1,126	3.02	8	0.71
Electricity, gas, steam and air conditioning supply	304	0.81	8	0.71
Financial and insurance activities	74	0.2	2	0.18
Human health and social work activities	359	0.96	2	0.18
Information and communication	1,421	3.81	31	2.74
Uncategorised manufacturing	5,984	16.04	147	13.00
Mining and quarrying	552	1.48	5	0.44
Other service activities	55	0.15	3	0.27
Professional, scientific and technical activities	2,752	7.38	84	7.43
Public administration and defence; compulsory social security	77	0.21	1	0.09
Real estate activities	10	0.03	0	0
Transportation and storage	7,164	19.2	209	18.48
Water supply; sewerage, waste management and re-mediation	247	0.66	10	0.88
Wholesale and retail trade; repair of motor vehicles and motorcycles	55	0.15	3	0.27
No sector declared	337	0.9	0	0
Total	37,308	100	1,131	100

Figure 2.D.1: Afghan formal sector's growth



Note: Authors calculation. Data are retrieved from the register of formal companies of the Afghan Ministry of Economy.

Figure 2.D.2: Altitude and major cities



Note: Altitude is the district average elevation in meters.



Table 2.D.3: Conflict exposure, additional specifications (1)

Specification:	IPW		Time het.		Coop. vs ext.	
	standard*	with sectors	elasticities	odds ratios	elasticities	odds ratios
	(1)	(2)	(3)	(4)	(5)	(6)
Military events	0.264 <sup>a</sup> (0.063)	0.305 <sup>a</sup> (0.066)	0.173 <sup>a</sup> (0.065)	0.042 (0.070)		
Mil. ev. shocks			-0.209 <sup>a</sup> (0.018)	-0.405 <sup>a</sup> (0.059)		
Mil. ev. shocks x Military events				0.053 <sup>a</sup> (0.016)		
Cooperative ev.					0.219 <sup>a</sup> (0.057)	1.551 <sup>a</sup> (0.141)
External ev.					-0.365 (0.233)	0.482 <sup>b</sup> (0.160)
Year FE	yes	yes	yes	yes	yes	yes
Firm FE	yes	yes	yes	yes	yes	yes
No. observ.	478.870	477.148	479.150	479.150	479.150	479.150

FE, Fixed effects; robust standard errors clustered at the firm level are in parentheses; standards errors of odds ratios are clustered at the district level. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively. Dependent variable is the firm's formal existence. Estimator is a logit. \*The standard IPW weights firms depending on their size and nationality

Table 2.D.4: Conflict exposure, additional specifications (2)

Specification:	Events' intensity (num. deaths)		Concentration		Provincial ev.	
	no distinction	soldiers vs civilians	spatial weights	temporal weights	elasticities	odds ratios
	(1)	(2)	(3)	(4)	(5)	(6)
Military events					0.370 <sup>a</sup> (0.063)	2.096 <sup>a</sup> (0.384)
Deaths	0.223 <sup>a</sup> (0.060)					
Deaths soldiers		0.342 <sup>a</sup> (0.075)				
Deaths civilians		-0.073 <sup>b</sup> (0.037)				
Weighted military events			-0.108 <sup>b</sup> (0.049)	-0.048 <sup>b</sup> (0.023)		
Provincial mil. ev.					-0.129 <sup>a</sup> (0.018)	0.772 <sup>a</sup> (0.047)
Year FE	yes	yes	yes	yes	yes	yes
Firm FE	yes	yes	yes	yes	yes	yes
No. observ.	420.947	417.337	479.150	479.150	479.150	479.150

FE, Fixed effects; robust standard errors clustered at the firm level are in parentheses; standards errors of odds ratios are clustered at the district level. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively. Dependent variable is the firm's formal existence. Estimator is a logit. In columns 3 and 4, events' weights depend on their distance from the average district events' location/date. Provincial military events exclude events in the district of interest

Table 2.D.5: IV and altitude non-linearity

Dependent variable:	Military events
Estimator:	OLS
	(1)
$\ln(\text{Mil.ev.out}) * \frac{\text{altitude}}{\text{altitude}}$	-0.000 (0.000)
$\ln(\text{Mil.ev.out}) * \frac{\text{altitude}}{\text{altitude}} * \text{Altitude}_{Q1}$	0.002 <sup>c</sup> (0.001)
$\ln(\text{Mil.ev.out}) * \frac{\text{altitude}}{\text{altitude}} * \text{Altitude}_{Q2}$	0.002 <sup>a</sup> (0.001)
$\ln(\text{Mil.ev.out}) * \frac{\text{altitude}}{\text{altitude}} * \text{Altitude}_{Q3}$	0.002 <sup>a</sup> (0.000)
Year FE	yes
Firm FE	yes
No. observ.	373,080

OLS Ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; stage standard errors are clustered at the district-year level. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 2.D.6: IV and NATO disengagement

Dependent variable:	Military events	
Estimator:	OLS	
	(1)	(2)
IV	1.023 <sup>a</sup> (0.255)	1.025 <sup>a</sup> (0.257)
IV*I(year>2012)	-0.172 <sup>a</sup> (0.055)	
IV*I(2013)		-0.170 <sup>a</sup> (0.060)
IV*I(2014)		-0.166 <sup>a</sup> (0.069)
IV*I(2015)		-0.170 <sup>a</sup> (0.080)
IV*I(2016)		-0.182 <sup>a</sup> (0.092)
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	522,312	522,312

OLS Ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; stage standard errors are clustered at the district-year level. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 2.D.7: IV, control function approach

Dependent variable: Firm's formal existence	
Second stage	(1)
Estimator:	Logit
Military events	4.264 <sup>a</sup> (0.963)
First stage residuals	-4.679 <sup>a</sup> (0.984)
Year FE	yes
Firm FE	yes
No. observ.	233.590
First-stage	
Estimator: OLS	
IV	1.025 <sup>a</sup> (0.240)
Year FE	yes
Firm FE	yes
No. observ.	373.080
KPW rk F-stat:	18
KPW rk LM-stat:	8

OLS Ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; first stage standard errors are clustered at the district-year level; second stage standard errors are clustered at the firm level. In second stage, the logit-FE estimator excludes observations with constant outcome – in our case firms, that always or never exist from 2003 to 2012. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Table 2.D.8: IV estimation, Firms heterogeneity (full table)

Dependent variable:	Firm's formal existence	
Second-stage	(1)	(2)
Estimator:	Logit	Logit
Military events	5.489 <sup>a</sup> (0.922)	5.457 <sup>a</sup> (0.928)
Military events x ln(size)	-0.411 <sup>a</sup> (0.048)	-0.458 <sup>a</sup> (0.048)
Military events x NATO		1.156 <sup>c</sup> (0.629)
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	115.580	115.420
First-stage		
Instrumented variable:	Mil. int.	
IV	1.063 <sup>a</sup> (0.237)	1.060 <sup>a</sup> (0.237)
IV x ln(size)	0.008 <sup>c</sup> (0.005)	0.008 <sup>c</sup> (0.005)
IV x NATO		0.003 (0.011)
SW F-stat:	37	61
SW Chi-sq	41	69
Instrumented variable: Mil. int. x ln(size)		
IV	0.103 (0.842)	0.093 (0.841)
IV x ln(size)	1.074 <sup>a</sup> (0.137)	1.074 <sup>a</sup> (0.136)
IV x NATO		0.007 (0.035)
SW F-stat:	100	132
SW Chi-sq	112	146
Instrumented variable: Mil. int. x NATO		
IV		0.002 (0.039)
IV x ln(size)		0.000 (0.001)
IV x NATO		1.052 <sup>a</sup> (0.153)
SW F-stat:		79
SW Chi-sq		88
Year FE	yes	yes
Firm FE	yes	yes
No. observ.	186.950	186.190

OLS, ordinary least square; FE, Fixed effects; robust standard errors are in parentheses; first stage standard errors are clustered at the district-year level; second stage standard errors are clustered at the firm level; median size is equal to 6. In first stages, only the coefficient of the interaction variable of interest is reported. Full table is available in appendices. Clustering the second-stage standard errors at the district level does not affect results. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

## Chapter 3

# Over-Distorted Gravity

# Welfare Gains from Trade and Bilateral Elasticities

We develop a theoretical framework for international trade and perform a welfare analysis in the presence of exporter-importer-specific trade elasticities. Adapting the workhorse models of trade, we show that in the presence of bilateral trade elasticities, the share of domestic expenditure and the elasticity of exports to the importer's price-index are sufficient statistics to infer welfare gains from trade. We provide a methodology for policy evaluation in general equilibrium. Applying this method to Regional Trade Agreements, we show that bilateral trade elasticities imply important distortions, which, if ignored, can lead to important misestimations of the trade liberalization's impact.

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What are the welfare gains from trade? This question at the core of the international economics literature has shaped decades of research. In their seminal paper, [Arkolakis et al. \(2012\)](#) show that modern models used to analyse gains from trade rely on the same conditions and have the same welfare implications. Based on these theoretical foundations, the literature developed a flexible policy evaluation approach, called initially the *calibrated share form* ([Dekle et al., 2007](#)) or, more recently, the *exact hat algebra*, leading to numerous extensions and applications ([Costinot and Rodríguez-Clare, 2014](#)).<sup>1</sup>

In this class of models, the trade elasticity is the key parameter. It determines how any change in trade costs impacts exports and real-incomes. In standard trade models, this parameter is a constant derived either from the elasticity of substitution (Armington model) or from the productivity distribution (Models with heterogeneous firms or Ricardian models) ([Arkolakis et al., 2012](#)). Recent literature actively discusses the trade elasticity, questioning its theoretical foundations, estimations and consequences. [Brooks and Pujolas \(2019\)](#) develop a model in which the trade elasticity is country-specific, leading to estimated gains from trade 25% higher than with constant elasticity on average. [Ossa \(2015\)](#) shows the importance of accounting for cross-industry variations in trade elasticity, while [Fontagné et al. \(2022\)](#) pursue the investigation at the product-level. [Bas et al. \(2017\)](#) analyse the role of the trade elasticity in heterogeneous firms models. They provide a methodology to quantify bilateral aggregate elasticities when gravity does not hold. [Mrázová and Neary \(2017\)](#) develop a large theoretical framework showing the different forms structural gravity can take depending on the assumed utility function. Interestingly, they exhibit that, for a set of utility functions, theoretical trade elasticities are country-pair-specific and negatively correlated with the value of bilateral trade. In related works, [Carrière et al. \(2020\)](#) and [Chen and Novy \(2022\)](#) provide empirical evidence of bilateral trade elasticities' existence consistent with modern structural gravity estimations. This empirical finding of bilateral elasticities raises the issue of understanding how gains from trade statistics derived in [Arkolakis et al. \(2012\)](#) must be adopted.

In this chapter, we address this challenge and fill this gap in the literature. We adapt the workhorse trade models to bilateral trade elasticities.<sup>2</sup> From this new theoretical framework, we derive the gains from trade and show that the elasticity of exports to the importer's price-index and the share of domestic expenditures are sufficient statistics. Furthermore, we provide a methodology to perform trade policy evaluations with bilateral trade elasticities.

The chapter is organised as follows. In section 1, we adapt the two categories of model on which

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<sup>1</sup>Some examples of papers based on this approach: [Anderson and Yotov \(2016\)](#); [Anderson et al. \(2018\)](#); [Allen and Arkolakis \(2022\)](#); [Caliendo et al. \(2019\)](#); [Dhingra et al. \(2017\)](#); [Fontagné and Santoni \(2021\)](#); [Mattoo et al. \(2022\)](#); [Mayer et al. \(2019\)](#); [Vandenbussche et al. \(2022\)](#).

<sup>2</sup>In the body of the text we present adaptations of the Armington model and CES monopolistic competition models with heterogeneous firms. In appendix, we propose an extension to the Ricardian models.

relies modern trade welfare analysis: Armington models ([Anderson, 1979](#); [Anderson and Van Wincoop, 2003](#); [Armington, 1969](#)) and heterogeneous firms models ([Chaney, 2008](#); [Melitz, 2003](#)). Introducing bilateral trade elasticities, we discuss the changes in the micro-foundation assumptions, their theoretical meanings and the derived structural gravity equations.

In section 2, we show that in the presence of bilateral trade elasticities, the share of domestic expenditure and the elasticity of exports to the importer's price-index, we call price-index elasticity, are sufficient to infer the welfare gains from trade. The price-index elasticity is importer-specific and derived from bilateral trade elasticities. Using a Method of Moments Quantile Regression, we estimate them. Empirically, bilateral trade elasticities decrease with the value of trade, while the price-index elasticity decreases with the importer economic size. Consequently, computing welfare gains from trade, we show that constant trade elasticities approach over-evaluate small economies' gains and under-evaluate large economies' ones. At the sector level, we can observe important heterogeneities between differentiated or homogeneous products fitting our theoretical and empirical methods.

In section 3, we develop a methodology to perform trade policy evaluations with bilateral trade elasticities. Using the Poisson Pseudo Maximum Likelihood (PPML) estimator properties, we adapt [Anderson and Yotov \(2016\)](#) methodology and its extension in [Anderson et al. \(2018\)](#). Three distortions lead to welfare results different from the constant trade elasticities approach: (i) because trade elasticities are exporter-importer-specific, a common trade cost shock will not have identical effects, (ii) country-specific elasticities of exports to importer and exporter price indexes induce heterogeneous change in aggregated prices and wages, (iii) the elasticity of exports to wage is also exporter-importer-specific, which distorts the negative effect of wages on trade. Using our method to evaluate Regional Trade Agreements' (RTAs) impact, we show that, in some circumstances, the constant trade elasticity approach over-estimates welfare gains.

Finally, we bring a short conclusion in section 4 discussing the main theoretical and empirical implications of the bilateral trade elasticities inclusion in welfare analysis. Accounting for bilateral trade elasticities in welfare analysis has important consequences. Therefore, omitting this bilateral dimension can lead to significant misestimations of the trade liberalization's impact.

### 3.1 Bilateral trade elasticities in theory

As shown by [Arkolakis et al. \(2012\)](#), modern analysis of welfare gains from trade are derived from a family of models sharing common conditions, including Armington models ([Anderson, 1979](#); [Anderson and Van Wincoop, 2003](#); [Armington, 1969](#)) and heterogeneous firms models ([Chaney, 2008](#); [Melitz, 2003](#)). In this section, we adapt these workhorse models of trade to bilateral trade elasticities and

propose economic interpretations of these theoretical changes. Besides, we present in the appendix section 3.A models' full derivation and propose in 3.B an extension to Ricardian models (Eaton and Kortum, 2002).

### 3.1.1 Trade elasticities design

First of all, we properly define the design of the bilateral trade elasticities which applies to all models. Bilateral trade elasticities are exporter-importer-specific. In other words, the trade elasticity is specific to the exporter-importer dyad and depend on how elastic exports are to a change in trade cost from a specific origin and toward a specific destination. Indeed, a given change in export cost in the countrypairs  $ij$  or  $lk$  may imply different reactions among exporters and consumers.

We define bilateral trade elasticities as the product of two parameters  $\theta$  and  $\zeta$ .<sup>3</sup>  $\theta$  is the main parameter. It reflects the general sensitivity of an economy to prices.  $\zeta$  is a deviation from  $\theta$ <sup>4</sup> To meet empirical evidence from Carrère et al. (2020); Chen and Novy (2022), we assume that  $\theta$  and  $\zeta$  are negatively correlated with the economic size of partners, but it is not strictly required for purely theoretical needs. The economic meaning of  $\theta$  and  $\zeta$  may depend on the theoretical framework. Therefore, we further discuss these parameters in each model that we present below.

### 3.1.2 Armington model

In the standard Armington model (Anderson, 1979; Armington, 1969), the trade elasticity is directly derived from the CES parameter. Hence, the only assumption that needs to be adapted is the form of the utility function:

$$U = \int_{\Omega} q(\omega)^{\frac{\theta_i \zeta_j - 1}{\theta_i \zeta_j}} d\omega \quad (3.1)$$

where  $q(\omega)$  denotes the quantity consumed of good  $\omega$ . Here, the only, but major, difference with the CES case is the presence of country-pair elasticities of substitution.<sup>5</sup> This means that consumers do not have the same willingness to switch from one variety to another depending on their location  $j$  and the product's origin  $i$ . Assuming  $\frac{\delta \theta_i}{\delta \frac{Y_i}{\Pi_i^{1-\theta_i}}} < 0$  and  $\frac{\delta \zeta_j}{\delta \frac{X_j}{\Phi_j^{1-\theta_j}}} < 0$ , with  $X_j$  the total expenditures in country  $j$ ,  $Y_i$  the total output in  $i$ ,  $\Phi_j$  the importer's price index and  $\Pi_i$  the exporter's price index, implies that

<sup>3</sup>Alternatively, we may write  $\theta_{ij}$ . Yet, as we detail throughout the chapter, using the product form i) facilitates the interpretation and makes visible the connection with the aggregate price elasticity, and ii) it formalizes the contribution of each partner to the elasticity which we need to compute in the empirical application.

<sup>4</sup>This means that we assume that an exporter (or respectively an importer) affects trade elasticity similarly to all of its partners. Whether  $\theta$  and  $\zeta$ , are, in the trade elasticity, exporter ( $i$ ) or importer ( $j$ ) specific may depend on the fundamental assumptions of theoretical models. Precisely, in our adaptations of the Armington model and the CESmonopolistic-competition models with heterogeneous firms, the main parameter ( $\theta$ ) is exporter-specific and the distortion parameter ( $\zeta$ ) is importer-specific. On the contrary, in the the Ricardian model adaptation,  $\theta$  is importer-specific and  $\zeta$  exporter-specific.

<sup>5</sup>As discussed by Carrère et al. (2020),  $U$  and  $\bar{U} = (U)^{\frac{\theta}{\theta-1}}$  have identical behaviour since  $U$  is a monotonically increasing transformation of  $\bar{U}$ , and preferences are ordinal.  $\bar{U}$  is a more familiar way of writing CES inspired utility function, yet for simplicity we use  $U$  in the model.



this willingness decreases with the importer's and the exporter's size. Such a hypothesis describes a "subconvex" demand, which is equivalent to the "Marshall's Second Law of Demand", assuming that the elasticity of demand decreases with sales (Mrázová and Neary, 2017; Carrère et al., 2020).<sup>6</sup> This means that: i) consumers have a preference for varieties that are rare in the global economy – i.e. from small origins  $i$ , ii) consumers in small destination economies  $j$  have a lower taste for variety.

This change in the utility function has immediate consequences on the relative demand function for varieties:<sup>7</sup>

$$\frac{q_{ij}}{q_{lj}} = \frac{\left(\frac{\theta_i \zeta_j - 1}{\theta_i \zeta_j} \lambda p_{ij}\right)^{-\theta_i \zeta_j}}{\left(\frac{\theta_l \zeta_j - 1}{\theta_l \zeta_j} \lambda p_{lj}\right)^{-\theta_l \zeta_j}} \quad (3.2)$$

or

$$q_{ij} = \left(\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \lambda p_{ij}\right)^{-\theta_i \zeta_j} X_j \int_l \left[ \left(\frac{\theta_l \zeta_j}{\theta_l \zeta_j - 1} \lambda\right)^{-\theta_l \zeta_j} p_{lj}^{1-\theta_l \zeta_j} \right] \quad (3.3)$$

Demands functions now depend on  $X_j$  the total expenditures in country  $j$ ,  $p_{ij}$  the price of the variety of origin  $i$  in the destination economy  $j$ , and  $p_{lj}$  with  $l$  an origin different from  $i$ . Besides the evident switch from constant price elasticities to bilateral ones, two new components appear.  $\lambda$ , derived from the Lagrangian, is interpretable as a preference parameter. Constant and exogenous, it affects quantities only through origin-destination differences in elasticities of substitution. In the presence of bilateral elasticities of substitution, mark-ups  $\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1}$  also affect relative quantities outside prices. Thus, at the country-pair level we observe:

$$X_{ij} = X_j \left(\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \lambda\right)^{-\theta_i \zeta_j} p_{ij}^{1-\theta_i \zeta_j} \Phi_j^{\theta_j - 1} \quad (3.4)$$

with the Dixit-Stiglitz price index:

$$\Phi_j = \left[ \int_l \left(\frac{\theta_l \zeta_j}{\theta_l \zeta_j - 1} \lambda\right)^{-\theta_l \zeta_j} p_{lj}^{1-\theta_l \zeta_j} \right]^{\frac{1}{1-\theta_j}} \quad (3.5)$$

And knowing that  $p_{ij} = \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} w_i \tau_{ij}$ , we obtain the gravity equation:

$$X_{ij} = X_j \frac{\left(\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1}\right)^{1-2\theta_i \zeta_j} \lambda^{-\theta_i \zeta_j} w_i^{1-\theta_i \zeta_j} \tau_{ij}^{1-\theta_i \zeta_j}}{\Phi_j^{1-\theta_j}} \quad (3.6)$$

<sup>6</sup>Carrère et al. (2020) and Chen and Novy (2022) proposed alternative models with bilateral trade elasticities based on non-CES demand, relying respectively on an additively separable preferences and a translog expenditure function. In comparison, our model lead to similar implications and partial equilibrium structure, but preserve tractability, flexibility and conditions to perform general-equilibrium welfare analysis as described by Arkolakis et al. (2012).

<sup>7</sup>See the appendix section 3.A.1 for the full model derivation.

or making visible output  $Y_i$  and outward multilateral resistances  $\Pi_i$ :<sup>8</sup>

$$X_{ij} = \frac{Y_i}{\Pi_i^{1-\theta_i}} \frac{X_j}{\Phi_j^{1-\theta_j}} \left( \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \right)^{1-2\theta_i \zeta_j} \lambda^{\theta_i(1-\zeta_j)} w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{1-\theta_j \zeta_i} \quad (3.7)$$

Doing so, we immediately observe three points. i) The trade elasticity – i.e. the elasticity of  $X_{ij}$  to  $\tau_{ij}$  – is now bilateral  $(1 - \theta_i \zeta_j)$ . ii) Multilateral resistances' elasticities are country-specific – importer-specific  $(1 - \theta_j)$  for the inward multilateral resistance  $\Phi_j$  and exporter-specific  $(1 - \theta_i)$  for the outward one  $\Pi_i$ . Therefore,  $\theta$ , the main parameter of the trade elasticity, reflects the general sensitivity of exports to prices. iii) The wage elasticity is equivalent to the trade elasticity (see eq. 3.6). Making visible in eq. (3.7) the output-outward multilateral resistance ratio  $\frac{Y_i}{\Pi_i^{1-\theta_i}}$ , the overall negative effect of wages on exports toward all destinations ( $w^{1-\theta_i}$ ) is captured. Yet, there subsists a bilateral distortion ( $w^{\theta_i(1-\zeta_j)}$ ). The consequences of this distortion on trade strongly depend on the value of  $\zeta_j$ , capturing in this model the destination-specific taste for variety. If  $\zeta_j > 1$ ,  $\theta_i(1 - \zeta_j)$  is negative and the effect of wages on exports toward  $j$  is amplified. On the contrary, if  $\zeta_j < 1$ ,  $\theta_i(1 - \zeta_j)$  is positive and the negative effect of wages on exports toward  $j$  is mitigated.<sup>9</sup>

Because in the presence of bilateral elasticities of substitution markups  $\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1}$  depend on origin  $i$  and destination  $j$ , they impact relative prices and quantities (eq. 3.2) and, therefore, exports (eq. 3.6). Furthermore, assuming  $\frac{\delta \theta_i}{\delta \frac{Y_i}{\Pi_i^{1-\theta_i}}} < 0$  and  $\frac{\delta \zeta_j}{\delta \frac{X_j}{\Phi_j^{1-\theta_j}}} < 0$  means that firms from a large economy  $i$  selling in a large destination market  $j$  have higher markup than firms from a small economy  $l$  selling in a small destination market  $k$ . Similarly, bilateral elasticities of substitution induce that preferences, captured by  $\lambda^{-\theta_i \zeta_j}$ , affect relative quantities (eq. 3.2) and exports (eq. 3.6).<sup>10</sup>  $\lambda$  is a constant and exogenous, it affects trade only through the differences in consumers' preferences that exporter-importer-specific elasticities of substitution  $\theta_i \zeta_j$  imply (eq. 3.1). Yet, as detailed in section 3.2,  $\lambda$  and  $\frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1}$  have no effect on welfare gains from trade liberalization.

### 3.1.3 Heterogeneous firms model

We now turn to the case of monopolistic competition with firm heterogeneity à la Melitz (2003).<sup>11</sup> We keep using the CES utility function: consumers in each country maximise their utility by consuming  $q(\omega)$  units of each differentiated goods  $\omega$  following  $U = [\int_{\Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} d\omega]^{\frac{\sigma}{\sigma-1}}$ . Assuming that the distribution of firms' productivity ( $\gamma$ ) follows a Pareto distribution as in Chaney (2008) and Helpman et al. (2008), the trade elasticity is immediately derived from the shape parameter of the Pareto:  $\theta$ . In

<sup>8</sup>With  $Y_i \Pi_i^{\theta_i - 1} = \lambda^{-\theta_i} w_i^{1-\theta_i}$  derived from the market clearance:  $Y_i = \sum_j X_{ij} = \sum_j X_j \left( \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \right)^{-\theta_i \zeta_j} p_{ij}^{1-\theta_i \zeta_j} \Phi_j^{\theta_j - 1} = \lambda^{-\theta_i} w_i^{1-\theta_i} \Pi_i^{1-\theta_i}$ .

<sup>9</sup>We further discuss the value of  $\zeta_j$  in section 3.3 where we empirically estimate it.

<sup>10</sup>As for wages, preferences effect is partially captured by the output-inward multilateral resistance ratio.

<sup>11</sup>See the appendix section 3.A.2 for full model derivation.

this framework, technology is constant: all countries' productivity distribution is the same. We now relax this assumption, allowing the Pareto shape parameter to be country-pair-specific, depending on an exporter-specific component  $\theta_i$  and an importer-specific component  $\zeta_j$ . Namely, we assume that marginal cost  $\alpha = \frac{1}{\gamma}$  follows the inverse Pareto distribution:

$$G_{ij}(\alpha) = (\alpha^{(\theta_i \zeta_j)} - \underline{\alpha}^{(\theta_i \zeta_j)}) / \overline{\alpha'}^{(\theta_i \zeta_j)} - \underline{\alpha}^{(\theta_i \zeta_j)} = \frac{\alpha^{(\theta_i \zeta_j)}}{\overline{\alpha'}^{(\theta_i \zeta_j)}} \quad (3.8)$$

with  $\overline{\alpha'} = \frac{X_j}{\Phi_j} \bar{\alpha}$  the maximum marginal cost<sup>12</sup> and  $\underline{\alpha}$  the minimum one.<sup>13</sup> Origin dependant  $\theta_i$  simply means that the technological level in countries may be different (Falvey et al., 2006). On the other hand, the destination component  $\zeta_j$  introduces sub-varieties in the model. Each firm still has a monopole on its variety. Yet, because of destination market restrictions or preferences, firms must adapt their production process, ending with a sub-variety sold only in its specific destination market.<sup>14</sup> This distorts the heterogeneity in productivity for sub-varieties specific to the destination market  $j$ , that  $\zeta_j$  captures.<sup>15</sup>

Matching with the stylized fact of smaller trade elasticities for larger countries, that means technology has been developed so that, on average, firms are more efficient where the real revenue is the highest. In this case, the destination-specific technological heterogeneity is interpretable as the global cost to adapt a technology designed to match the restrictions of dominant markets to smaller market restrictions.

Therefore, each country-pair faces a specific distribution of productivity, while each firm owns different productivities depending on the destination market they are producing for. Since each country-pair productivity is Pareto distributed, the observed average productivity of firms in a given country is also Pareto distributed.<sup>16</sup> Furthermore, in this model, the price and profit equations are very similar to

<sup>12</sup>Assuming  $\overline{\alpha'} = \frac{X_j}{\Phi_j} \bar{\alpha}$ , we amplify the dependence of the productivity on the destination market. The intuition is similar to  $\zeta_j$ : firms adapt their productivity to be the most efficient on the largest markets. The main consequence is that the impact of demand on firms' entry is not influenced by the distribution of productivity (see in appendix section 3.A.2 for details). This assumption is not required to obtain bilateral trade elasticities, but better matches with the empirical evidences presented by Carrère et al. (2020) and Chen and Novy (2022) and facilitates comparison with the Armington framework. Removing this assumption lead to an origin-specific elasticity of expenditures ( $X_j$ ) in the gravity equation that must be considered in the general equilibrium computation described in section 3.3.

<sup>13</sup>For solving issues we follow Helpman et al. (2008) and assume that  $\underline{\alpha} = 0$ , i.e. that there is always at least one firm productive enough to export even a very small value.

<sup>14</sup>For example, lets take a firm producing blue T-shirts. In all markets this firms sell blue T-shirts. Nonetheless, sizes of consumers in Japan and in the US are not the same which force the firm to adapt its production, like other firms producing T-shirts, to this restriction. Another striking example is the necessity for car-producing firms to adapt the wheel's location to sell in the UK and in France – such product differences won't be observe by the consumer, but need a production process adaptation from the firm.

<sup>15</sup>We focus on the simplest way of designing this destination-specific distortion in the productivity distribution. The latter implies that a firm may be poorly productive for the sub-variety sold in  $j$  but highly productive for the one sold in  $k$ . Alternatively, we can assume that firms do not draw a productivity for each destination but a position in the productivity distributions which applies to all sub-varieties. Thus, its productivity is still exporter-importer-specific but the firm can no longer have different positions in the productivity distributions depending on the destination. This does not affect our conclusions.

<sup>16</sup>At the country-pair level, many firms draw small productivity while a few firms draw a high one. Taking the firms' average productivity, we observe the same thing. Many firms draw small productivity for all destinations while a few firms draw only high productivity. Nonetheless, in this framework, some firms that are on average not very productive

those described in Melitz (2003), the only change being the bilateral productivity parameter. Thus, deriving the model, we obtain the following expression of the structural gravity equation with bilateral trade costs elasticities:

$$X_{ij} = N_i \bar{\alpha}_i^{-(\theta_i \zeta_j)} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_i \zeta_j} w_i^{-(\theta_i \zeta_j)} \frac{X_j}{\Phi_j^{-\theta_j} \tau_{ij}} \tau_{ij}^{-\theta_i \zeta_j} f_{ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1}-1]} \quad (3.9)$$

with

$$\Phi_j = \left( \sum_l N_l \bar{\alpha}_l^{-\theta_l \zeta_j} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_l \zeta_j} w_l^{-\theta_l \zeta_j} \tau_{lj}^{-\theta_l \zeta_j} f_{lj}^{-[\frac{\theta_l \zeta_j}{\sigma-1}-1]} \right)^{-\frac{1}{\theta_j}} \quad (3.10)$$

or, knowing from the market clearance condition that  $\frac{Y_i}{\Pi_i^{-\theta_i}} = N_i \bar{\alpha}_i^{-\theta_i} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_i \zeta_j} w_i^{-\theta_i}$ :

$$X_{ij} = \frac{Y_i}{\Pi_i^{-\theta_i}} \frac{X_j}{\Phi_j^{-\theta_j} \bar{\alpha}_i^{-(\theta_i \zeta_j)}} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{\theta_i(1-\zeta_j)} w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{-\theta_i \zeta_j} f_{ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1}-1]} \quad (3.11)$$

Compared to the standard Chaney (2008) gravity equation, we observe three main changes: i) The trade elasticity  $-(\theta_i \zeta_j)$  is bilateral and composed of an exporter-specific component  $\theta_i$  and a destination-specific component  $\zeta_j$ . Furthermore, note that the fixed trade cost elasticity  $-\left[\frac{\theta_i \zeta_j}{\sigma-1}-1\right]$  is also exporter-importer-specific. ii) The elasticity of exports to price-indexes ( $\Phi_j$  and  $\Pi_i$ ) is country-specific. These elasticities are defined by  $\theta$ , capturing the technological level of the country, meaning that the trade of an economy with concentrated weak productivities is more sensitive to variations in prices. iii) In this framework, the elasticity of export to wage is also equivalent to the trade elasticity (see eq. 3.10). Switching from eq. (3.10) to eq. (3.11), the output-price-index ratio  $\frac{Y_i}{\Pi_i^{-\theta_i}}$  captures the overall negative effect of wages on exports, but there subsists again a bilateral distortion  $w_i^{\theta_i(1-\zeta_j)}$ . If  $\zeta_j > 1$ , the distortion increases the negative effect of the wage on trade, and if  $\zeta_j < 1$ , it mitigates this negative effect. In the heterogeneous firms framework, this means that the wage affects less negatively exports for destination-specific sub-varieties whose productivities are the least concentrated.

## 3.2 Gains from trade

### 3.2.1 Defining gains from trade

In this class of models with bilateral trade elasticities, we then derive the theoretical gains from trade. We can define changes in welfare  $W$  – i.e. in real revenue – as follows (proof in appendix section 3.C):

$$\hat{W}_j = \hat{\psi}_{jj}^{-\frac{1}{\theta_j}} \quad (3.12)$$

---

may be poorly productive for the large majority of destinations, but very productive for a few destinations.

Interestingly, even in the presence of bilateral trade elasticity, [Arkolakis et al. \(2012\)](#)'s sufficient statistic demonstration holds. The share of domestic expenditure ( $\psi_{jj}$ ) and the elasticity of exports to the importer price-index ( $\theta_j$ ) – we call price-index elasticity – are sufficient to infer welfare changes in country  $j$ .

In [Arkolakis et al. \(2012\)](#) framework, the trade elasticity and the price-index elasticity are assumed as equal. As we showed in our models, this is no longer the case in a bilateral trade elasticity framework: the trade elasticity is exporter-importer-specific and defined by  $\theta_i$  and  $\zeta_j$ , while the price-index elasticity is importer-specific and defined by  $\theta_j$ . The economic interpretation of  $\theta$  depends on models, but in the gravity equations with bilateral trade elasticity, it always captures the sensitivity of exports to changes in aggregated prices.

Deriving the theoretical welfare gains from trade (see appendix section [3.C](#)), we observe that this is no longer the trade elasticity which matters but  $\theta_j$ , the elasticity of exports to the importer's price index (eq. [3.12](#)). This change induces country-specific welfare sensitivity to the openness to trade captured by  $\psi_{jj}$ . Below, based on eq. [\(3.12\)](#), we present how this affects the estimation of welfare gains from trade.

### 3.2.2 Data

In the following empirical application of our theoretical framework, we use common data sources. Trade data are computed from [Fontagné and Santoni \(2021\)](#) dataset, adapting the Tradeprod base of the CEPII ([De Sousa et al., 2012](#); [Mayer et al., 2023](#)) to general equilibrium analyses.<sup>17</sup> It covers 111 countries over the period 1990-2014. We use MacMap tariffs data ([Guimbard et al., 2012](#)), available for 2001, 2004, 2007, 2010, 2013 and 2016. The initial base is at the HS6 level. In our setting we use the exporter-importer-year average of tariffs. Wage data are retrieved from the International Labour Organisation database, available from 2011 to 2022. Country coverage may change depending on the year. We use the country-year average wage in US dollar. Finally, information on Regional Trade Agreements (RTAs) as well as other standard gravity variables, such as distance or language, are extracted from the Cepii Gravity database ([Conte et al., 2022](#)).

In our broadest panel, we follow 12321 exporter-importer pairs from 1990 to 2014. Including tariffs, the period is restricted to the years 2001, 2004, 2007, 2010 and 2013. To overcome this important fall in the sample size and observable variations, we extend our trade data with the BACI database from the CEPII ([Gaulier and Zignago, 2010](#)) when using tariffs. This increases the number of country-pairs to 25900 and allows us to include the year 2016.<sup>18</sup> When wages data are necessary, the panel is also

<sup>17</sup>Intranational flows are filled by linear interpolation of non-missing data, whereas the remaining missing values are extrapolated using country total exports ([Baier et al., 2019](#); [Fontagné and Santoni, 2021](#))

<sup>18</sup>The BACI database provide information at the HS6 level. Yet, in our empirical work we do not mobilise the sector-level variations. Therefore, trade data are aggregated at the country-pair-year level.

largely reduced. The number of country pairs is reduced to 4062 and the period limited to 2011-2014.

### 3.2.3 Estimating elasticities

To compute the welfare gains from trade, we need to estimate  $\theta_j$ . It can be directly retrieved from the bilateral trade elasticities  $\theta_i\zeta_j$ .  $\zeta$  being defined as a deviation from the main trade elasticity parameter  $\theta$  (see section 3.1.1), we can compute a good proxy of  $\theta_j$  using the importer's average bilateral trade elasticity. Yet, this first requires estimating those exporter-importer-specific trade elasticities.

To estimate the bilateral trade elasticities we need two elements: the average trade elasticity  $\bar{\theta}$ , and the exporter-importer-specific deviation from this average.

First we estimate the average trade elasticity. It can be derived from the tariffs' elasticity in a structural gravity estimation (Head and Mayer, 2014; Anderson et al., 2018). Thus we estimate:

$$X_{ijt} = \exp(-\bar{\theta}\ln(\text{tariff}s_{ijt} + 1) + \chi_{it} + \chi_{jt} + \chi_{ij}) * \varepsilon_{ijt} \quad (3.13)$$

where the exporter-year and importer-year fixed effects  $\chi_{it}$  and  $\chi_{jt}$  capture output, expenditures and multilateral resistances (see eq. 3.7 and eq. 3.11). The exporter-importer fixed-effects  $\chi_{ij}$  control for all time-invariant trade costs and parameters.<sup>19</sup> In this setting, we identify  $\bar{\theta}$  using the within variation of tariffs.<sup>20</sup> We use a Poisson Pseudo Maximum Likelihood (PPML) estimator.<sup>21</sup> In our sample, we find  $\bar{\theta} = 3.383$  (see appendix table 3.E.1, column 1).

Second, we estimate the deviation from the average trade elasticity. Based on our theoretical framework, this deviation is highly correlated with the value of trade. With such a constraint, the best solution is to use a Method of Moments-Quantile Regression (MMQR) estimator (Carrère et al., 2020). It is specifically designed to estimate the coefficients' deviation from the average depending on the value of the explained variable (Machado and Silva, 2019). Thus, we estimate:

$$X_{q,ijt} = -\widetilde{\theta}_i\zeta_j\ln(\text{tariff}s_{q,ijt} + 1) + BZ_{q,ijt} + \chi_{q,it} + \chi_{q,jt} + \varepsilon_{q,ijt} \quad (3.14)$$

where  $q$  is a defined quantile. The estimator being linear, we use the inverse hyperbolic sine transformation of  $X_{q,ijt}$ .<sup>22</sup>  $\chi_{q,it}$  and  $\chi_{q,jt}$  are the exporter-year and importer-year fixed effects. They capture

<sup>19</sup>Precisely, parameters captured by the exporter-importer fixed-effects fixed effects are:  $(\frac{\theta_i\zeta_j}{\theta_i\zeta_j-1})^{1-2\theta_i\zeta_j}$  and  $\lambda^{\theta_i(1-\zeta_j)}$  in the Armigton model, or  $\frac{\theta_i}{\alpha_i^{\theta_i(1-\zeta_j)}}$  and  $(\frac{\sigma-1}{\sigma})^{\theta_i(1-\zeta_j)}$  in the heterogeneous firms model.

<sup>20</sup>Estimating tariffs' elasticities by relying both on the cross-sectional and within variation of tariffs improves the precision of estimates (Fontagné et al., 2022). Yet, excluding exporter-importer fixed effects tends to strongly overestimate the average elasticity (Baier and Bergstrand, 2007; Heid et al., 2021). Therefore, we chose to keep these exporter-importer fixed effects in our estimation. We observe that, indeed, removing them leads to a higher estimate of the average trade elasticity.

<sup>21</sup>Gravity equation being non-linear, the PPML estimator allows us to take into account zero trade observations and avoid the biases caused by a combination of log-linearisation and heteroscedasticity (Silva and Tenreyro, 2006).

<sup>22</sup>Following Carrère et al. (2020), only intra-national trade flows are excluded.

the outputs, expenditures and multilateral resistances (see eq. 3.7 and eq. 3.11). The MMQR estimator does not allow for exporter-importer fixed effects.<sup>23</sup> Time-invariant trade costs are controlled through  $Z$ , a vector of time-invariant bilateral variables which may affect trade costs: the distance, the colonial past, the common language, the common currency, the common religion and the territorial contiguity.<sup>24</sup>

Following our theoretical framework, if the assumption that  $\theta$  and  $\zeta$  are negatively correlated with the economic size of partners is true, the MMQR estimates of  $-\widetilde{\theta}_i\zeta_j$  should increase with the percentile of exports. The estimated bilateral deviation of trade elasticities, with bootstrapped 95% confidence intervals, are graphically displayed in figure 3.2.1. The tariffs' coefficients are increasing in the value of trade, which confirms our assumption. From these results, we can derive the distribution of bilateral elasticities  $D(-\widetilde{\theta}_i\zeta_j)$  depending on the value of bilateral exports.

However, we need to correct this distribution so that its average matches the average trade elasticity estimated in equation (3.13). Therefore, we multiply each elasticity estimated with the MMQR by the ratio between the average trade elasticity from eq. (3.13) and the average elasticity in  $D(-\widetilde{\theta}_i\zeta_j)$  (with  $-\theta_i\zeta_j = -\widetilde{\theta}_i\zeta_j * \bar{\theta}/\bar{\theta}_D$ ).<sup>25</sup> Bilateral trade elasticities are displayed in figure 3.2.2. Finally, by using the country-specific average of bilateral trade elasticities, we can compute a good proxy of  $\theta_j$ , the elasticity of exports to the importer price-index (see eq. 3.7, 3.11 and 3.12).<sup>26</sup> In table 3.2.1, we report the estimated price-index elasticities, as well as their distortion compared to the average trade elasticity, which we would use in a constant trade elasticity framework (Anderson et al., 2018; Arkolakis et al., 2012; Costinot and Rodríguez-Clare, 2014). The observed distortion is clear. Small economies' elasticity increases, while large economies' elasticity decreases. In our sample, it is inbetween  $[-9.56\%; 6.41\%]$ ,  $-9.56\%$  being for the USA with  $\theta_j = 3.06$ , and  $6.41\%$  for Burundi with  $\theta_j = 3.60$ . Knowing that the price index  $\Phi_j = (\sum_i p_{ij}^{-\theta_i\zeta_j})^{-\frac{1}{\theta_j}}$  (see eq. 3.5 and 3.10), this means that small economies' price indexes are less sensitive to changes in prices than large economies.<sup>27</sup>

<sup>23</sup>Besides, trade cost related variables, some parameters are therefore only partially captured by fixed-effects (see footnote 19). Yet, they are all exogenous parameters, and, empirically, should not be correlated with aggregated tariffs. Our results should therefore not be biased. Furthermore, alternative estimators that allow for country-pair fixed effects show also a negative correlation between the value of bilateral trade and elasticities (Chen and Novy, 2022; Sotelo, 2019). The MMQR estimator allowing for important granularity, we nonetheless prefer this approach.

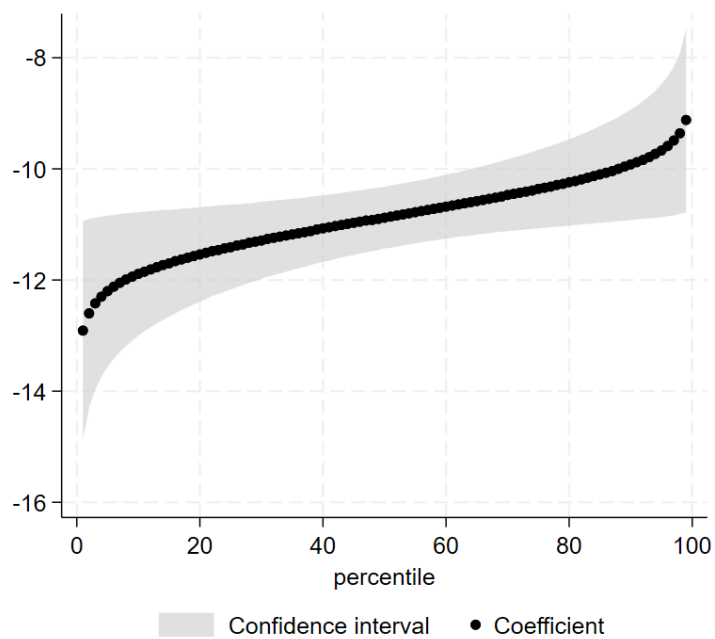
<sup>24</sup>In Carrère et al. (2020), the MMQR estimation is cross-section. In our case we also use the panel dimension to maximise statistical power. Performing a robustness check in cross-section, coefficients are unaffected.

<sup>25</sup>In our sample,  $\bar{\theta}/\bar{\theta}_D = 3.383/10.808$

<sup>26</sup>We compute  $\theta_j = \frac{\sum_i \theta_j \zeta_i}{I}$ . By assumption  $\zeta$ , the deviation parameter, is equal to one on average. This assumption is empirically verified. We observe that  $\bar{\zeta}$  is very close to one and  $\zeta \in ]0; 2[$  for all countries. Therefore we have:  $\frac{\sum_i \theta_j \zeta_i}{I} = \theta_j \bar{\zeta} = \theta_j$ .

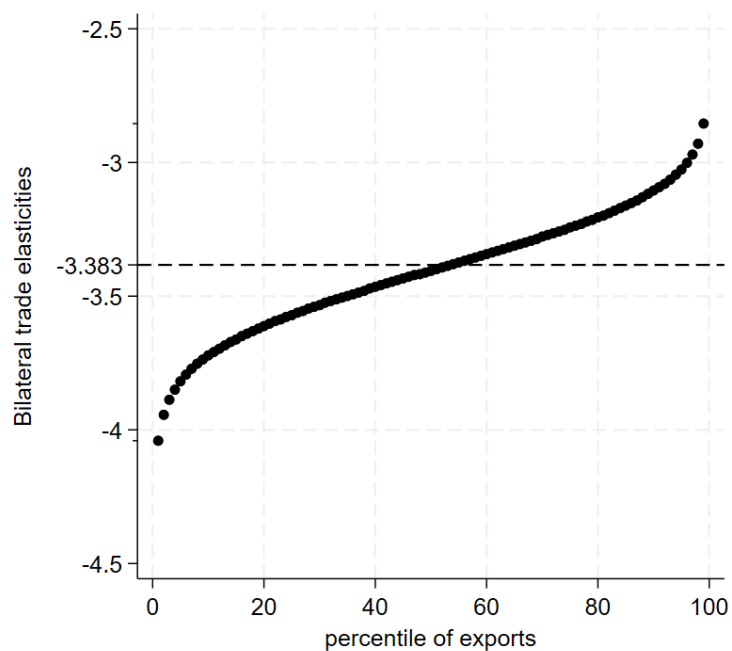
<sup>27</sup>With  $\Phi_j = (\sum_i p_{ij}^{-\theta_i\zeta_j})^{-\frac{1}{\theta_j}}$ , the higher  $\theta_j$ , the less a change in  $p_{ij}$  affects  $\Phi_j$ . See appendix section 3.C for more details.

Figure 3.2.1: Quantile estimates of tariffs



*Note:* Estimator is a Method of Moments-Quantile Regression with exporter-year and importer-year fixed effects.

Figure 3.2.2: Bilateral trade elasticities



*Note:* -3.383 is the average trade elasticity in our sample.



Table 3.2.1: Price index elasticities

Country(Iso3)	$\theta_j$	Change (in %) $(\frac{\theta_j}{\bar{\theta}} - 1) * 100$	Country(Iso3)	$\theta_j$ $(\frac{\theta_j}{\bar{\theta}} - 1) * 100$	Change (in %)
ALB	3.57	5.55	JOR	3.39	0.26
ARE	3.28	-3.15	JPN	3.07	-9.07
AUS	3.22	-4.79	KEN	3.40	0.58
AUT	3.19	-5.78	KOR	3.08	-8.80
BDI	3.84	13.49	KWT	3.40	0.48
BGR	3.30	0.76	LKA	3.33	-1.71
BRA	3.13	-5.13	MAR	3.31	-2.13
CAN	3.18	-6.00	MEX	3.21	-5.05
CHE	3.13	-7.40	MNG	3.70	9.25
CHL	3.29	-2.85	MUS	3.42	1.11
CHN	2.97	-12.08	MYS	3.15	-6.96
COL	3.33	-1.51	NER	3.62	7.21
DEU	3.04	-10.03	NOR	3.23	-4.40
DNK	3.19	-5.66	NZL	3.26	-3.63
DZA	3.49	3.08	OMN	3.38	-0.17
ECU	3.39	0.33	PER	3.32	-1.81
EGY	3.27	-3.32	PHL	3.29	-2.78
ESP	3.12	-7.88	POL	3.19	-5.62
FIN	3.21	-4.97	PRT	3.23	-4.38
FJI	3.66	8.29	QAT	3.34	-1.14
FRA	3.06	-9.30	ROM	3.36	-3.69
GBR	3.09	-8.64	SAU	3.31	-2.12
GRC	3.28	-3.17	SER	3.40	0.55
HUN	3.25	-3.84	SWE	3.16	-6.61
IDN	3.16	-6.47	TUN	3.35	-0.86
IND	3.08	-8.98	TUR	3.14	-7.10
IRL	3.20	-5.45	TZA	3.45	2.11
IRN	3.47	2.65	USA	3.01	-10.92
ISL	3.47	2.56	YEM	3.61	6.75
ISR	3.26	-3.55	ZAF	3.20	-5.49
ITA	3.08	-9.10	ZWE	3.64	7.69

Note: We estimated  $\bar{\theta} = 3.383$ .

## Sector level approach

Until then, we perform our analysis using country-pair level data. Yet, at the sector level, bilateral trade elasticities may follow different dynamics. Mobilising this source of variation may therefore lead to different results (Breinlich et al., 2022; Fontagné et al., 2022; Ossa, 2015). We reproduce the estimation of the bilateral deviation from the average trade elasticity with an MMQR estimator (eq. 3.14) at the HS1 sector level.<sup>28</sup> Results are displayed in figure 3.2.3. Despite a gain in precision thanks to the increase in the statistical power, the estimated deviation is very close to the one estimated with country-pair level data.<sup>29</sup>

Nevertheless, the theoretical interpretation of bilateral trade elasticities relying highly on the behaviour of consumers or firms regarding varieties and differentiation, the deviation from the average elasticity may strongly depend on whether the product is differentiated or not (Erkel-Rousse and Mirza, 2002; Feenstra et al., 2001; Rauch, 1999). To test this, we separately replicate the bilateral deviation estimation at the HS1 sector level for differentiated and homogeneous products.<sup>30</sup> MMQR results are presented in figure 3.2.4. Interestingly, for differentiated products, the deviation is very close to the baseline one. On the other hand, for homogeneous products the deviation is completely different and positively correlated with the value of bilateral trade. In our Armington theoretical framework, this new shape of bilateral trade elasticities is interpretable as a penalisation of differentiation when products are homogeneous: consumers have a taste for variety but value more the ones they consume the most.<sup>31</sup> Alternatively, in our framework with heterogeneous firms, the deviation specific to homogeneous products is interpretable as a stronger market-share and productivity concentration in large country-pair markets that may reflect economies of scale.<sup>32</sup>

In the rest of the chapter, we keep using aggregated data. Nonetheless, the flexibility of our theory and method allow for both country-pair and sector level approaches and any kind of bilateral trade elasticity.

### 3.2.4 Computing gains from trade

Once the bilateral trade and price-index elasticities are estimated, we can compute the gains from trade for the year 2013<sup>33</sup> using eq. (3.15). The share of domestic expenditures  $\psi_{jj}$  is directly observed from

<sup>28</sup>We keep using MacMap and BACI databases but aggregate at the HS1 sector level.

<sup>29</sup>The average trade elasticity may also be different depending on sectors. In this section we focus on the sector dependency of the bilateral heterogeneity of trade elasticities, but sector-specific average trade elasticities can be separately estimated and then considered using the method previously described. See (Fontagné et al., 2022) for more details on how to estimate sector-specific average trade elasticities.

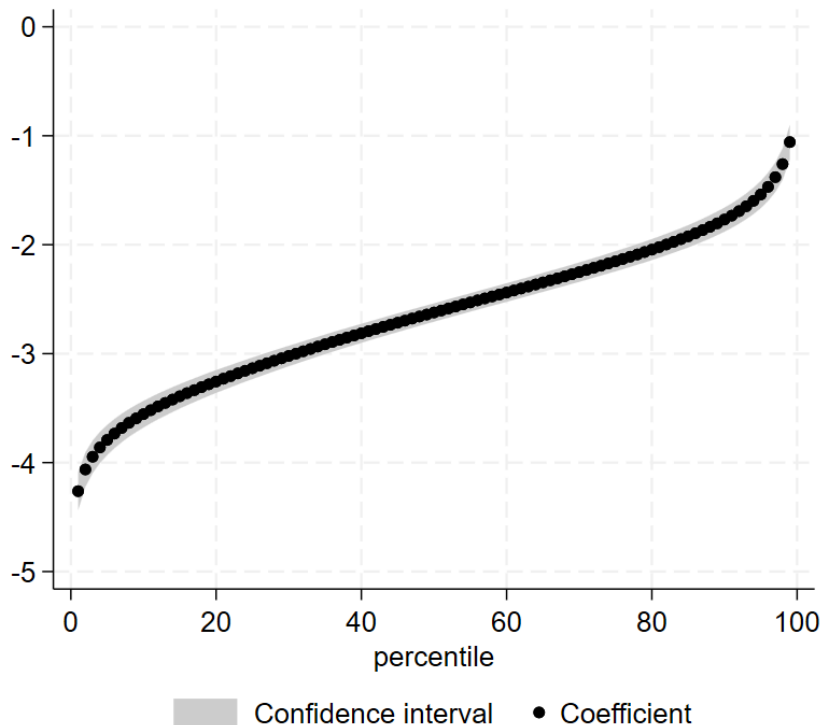
<sup>30</sup>Using Rauch (1999)'s classification, we include in homogeneous products both goods with a referenced price and goods traded on an organized exchange.

<sup>31</sup>See Mrázová and Neary (2017) for more details about *superconvex* demand.

<sup>32</sup>In the presence of economies of scale, the larger the market the more market-shares are concentrated in a few highly productive firms (Bain, 1954; Head and Ries, 2001). Assuming Pareto distribution of productivity, this means the shape parameter increases with the market's size (Chaney, 2008; Helpman et al., 2008).

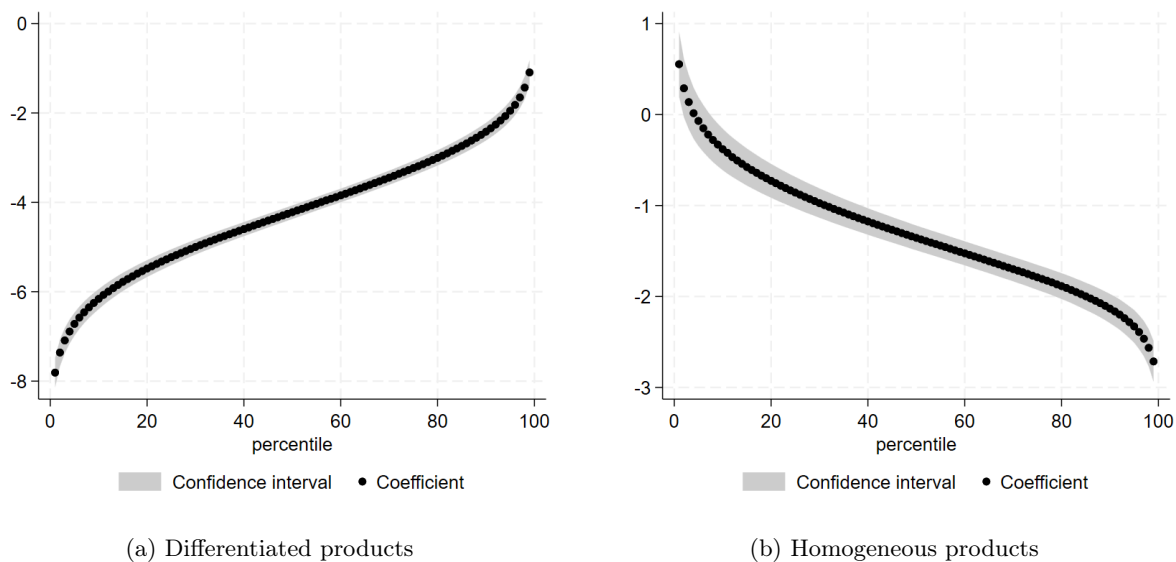
<sup>33</sup>The last year we observe trade, tariffs and wages in our sample.

Figure 3.2.3: Quantile estimates of tariffs, sector level data



*Note:* Estimator is a Method of Moments-Quantile Regression with exporter-year and importer-year fixed effects. Sector data are at the HS1 level.

Figure 3.2.4: Quantile estimates of tariffs, differentiated and homogeneous products



*Note:* Estimator is a Method of Moments-Quantile Regression with exporter-year and importer-year fixed effects. Sector data are at the HS1 level. Differentiated/homogeneous products are separated using Rauch (1999)'s classification.

data. Null observations are excluded.<sup>34</sup> We use the elasticities  $\theta_j$  from the previous section reported in table 3.2.1.

We report welfare results in table 3.2.2. Columns 1 & 5 present gains from trade with  $\theta_j$ , columns 2 & 6 with constant elasticity  $\theta = 3.383$ , and columns 3 & 7 changes in gains. Bilateral trade elasticities models induce noticeable changes in gains from trade. They are directly driven by the price-index elasticity  $\theta_j$ : a higher (lower) elasticity induces lower (higher) gains from trade. With a general perspective, these results show that constant trade elasticity approaches over-evaluate gains from trade for small economies like Burundi (by 10.73%), Yemen (by 5.75%) or Fiji (by 5.61%), and under-evaluates gains for large economies such as France (by 8.99%), the USA (by 11.60%) or China (by 13.58%).

### 3.3 Policy evaluation

From our theoretical models and the derivation of the gains from trade, we develop in this section a policy evaluation methodology. Working through our demonstration, we analyse the welfare effect of Regional Trade Agreements (RTAs).

#### 3.3.1 General equilibrium estimation

Depending on the chosen theoretical frame (Armington, Ricardian, or heterogeneous firms), gravity equation and price indexes are not exactly the same. Yet, by simplifying the General Equilibrium (GE) system of each model by dropping constant parameters which do not contribute to the effect of a change in trade costs,<sup>35</sup> we can define a common GE system with four equations:

$$X_{ij} = \frac{Y_i}{\Pi_i^{-\theta_i}} \frac{X_j}{\Phi_j^{-\theta_j}} w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{-\theta_i \zeta_j} \quad (3.15)$$

$$\Pi_i^{-\theta_i} = \sum_j \frac{w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{-\theta_i \zeta_j}}{\Phi_j^{-\theta_j}} X_j \quad (3.16)$$

$$\Phi_j^{-\theta_j} = \sum_i \frac{w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{-\zeta_j \theta_i}}{\Pi_i^{-\theta_i}} Y_i \quad (3.17)$$

$$w_i = \left( \frac{Y_i}{\Pi_i^{-\theta_i}} \right)^{-\frac{1}{\theta_i}} \quad (3.18)$$

<sup>34</sup> $\psi_{jj} = 0$  is inconsistent with the models. 48 countries are dropped.

<sup>35</sup>Although they are not specified, using the property of the PPML estimator to solve the GE system, those parameters are captured by the exporter-importer fixed effects (see footnote 19), which ensures we are consistent with the three models (Anderson and Yotov, 2016; Anderson et al., 2018).

Table 3.2.2: Gains from trade

Country (Iso3)	Gains (in %) constant elas.	Gains (in %) non-constant elas.	Change (in %) $(\frac{non-constant}{constant} - 1) * 100$	Country (Iso3)	Gains (in %) constant elas.	Gains (in %) non-constant elas.	Change (in %) $(\frac{non-constant}{constant} - 1) * 100$
ALB	54.20	52.28	-3.55	JOR	23.24	23.19	-0.23
ARE	79.51	80.54	1.30	JPN	5.92	6.49	9.65
AUS	14.26	14.92	4.63	KEN	17.71	17.62	-0.52
AUT	28.30	29.74	5.11	KOR	8.61	9.41	9.18
BDI	20.33	18.14	-10.73	KWT	15.37	15.30	-0.44
BGR	28.85	29.47	2.14	LKA	21.40	21.33	1.53
BRA	5.92	6.37	7.66	MAR	21.81	22.22	1.92
CAN	23.05	24.33	5.54	MEX	29.86	31.17	4.39
CHE	54.25	57.02	5.11	MNG	72.51	69.33	-4.38
CHL	54.58	55.62	1.91	MUS	42.39	42.03	-0.82
CHN	2.07	2.36	13.58	MYS	32.25	34.20	6.03
COL	18.96	19.23	1.38	NER	42.95	40.75	-5.11
DEU	16.71	18.39	10.06	NOR	22.77	23.68	4.01
DNK	28.30	29.72	5.01	NZL	17.12	17.70	3.41
DZA	13.64	13.25	-2.78	OMN	38.36	38.41	0.13
ECU	29.69	29.61	-0.28	PER	13.48	13.71	1.70
EGY	21.48	22.13	3.025	PHL	32.98	33.74	2.31
ESP	15.86	17.10	7.78	POL	19.74	20.79	5.30
FIN	15.33	16.06	4.79	PRT	22.65	23.55	3.99
FJI	47.17	44.52	-5.61	QAT	38.23	38.58	0.89
FRA	20.88	22.76	8.99	ROM	28.33	29.24	3.21
GBR	22.71	24.57	8.20	SAU	26.86	27.36	1.84
GRC	18.24	18.78	2.95	SER	19.29	19.19	-0.49
HUN	41.75	42.99	2.97	SWE	22.47	23.85	6.16
IDN	16.19	17.21	6.29	TUN	32.70	32.94	0.71
IND	7.33	8.026	9.46	TUR	12.88	13.80	7.09
IRL	39.27	40.99	4.39	TZA	42.85	42.19	-1.55
IRN	4.40	4.29	-2.53	USA	9.40	10.49	11.60
ISL	41.90	42.04	-1.88	YEM	18.03	16.99	-5.75
ISR	20.90	21.47	3.26	ZAF	16.53	17.40	5.27
ITA	12.48	12.72	9.34	ZWE	48.12	45.63	-5.17

Notes: The real revenue is our measure of welfare. Column 1 and 2 numbers are variations in percentage. Column 3's change displays  $(\frac{non-constant}{constant} - 1) * 100$ .

Equation (3.15) defines structural gravity.  $\Pi_i$  and  $\Phi_j$  are respectively the outward and inward multilateral resistance terms, while equation (3.18), derived from the market-clearance condition, defines the relation between wage and real output. We solve this welfare system by adapting the [Anderson and Yotov \(2016\)](#) methodology and its extension in [Anderson et al. \(2018\)](#) based on the PPML estimator properties.

First, we need to estimate the partial equilibrium effect of RTAs on trade. Using a PPML estimator, we regress the following equation:

$$X_{ijt} = \exp(\beta RTA_{ijt} + \theta_i(1 - \zeta_j)\ln(w_{it}) + \chi_{it} + \chi_{jt} + \chi_{ij}) * \epsilon_{ijt} \quad (3.19)$$

where  $RTA_{ijt}$  is a dummy variable taking the value one if exporter  $i$  and importer  $j$  have an RTA at time  $t$  and zero otherwise.  $\chi_{it}$  are exporter-year fixed effects,  $\chi_{jt}$  importer-year fixed effects and  $\chi_{ij}$  exporter-importer fixed effects.

To be consistent with the gravity equation (eq. 3.15), we must take into account that the wage elasticity  $\theta_i(1 - \zeta_j)$  is bilateral. Yet, the wage being exporter-time-specific, its variation is captured by the exporter-time fixed effect. Therefore, we need to replace the wage by an exporter-importer-time-specific variable which is still consistent with the GE system. The easiest solution is to include in the variable the destination-specific component of the wage elasticity – i.e. to substitute  $(1 - \zeta_j)\ln(w_{it})$ . This requires that we calibrate  $\zeta_j$ . We can retrieve a good proxy of  $\zeta_j$  from the bilateral elasticities  $(-\theta_i\zeta_j)$  that we estimated in section 3.2. From the theoretical definition of the bilateral trade elasticity, we know that  $-\zeta_j$  is equal to the bilateral trade elasticity  $(-\theta_i\zeta_j)$  divided by the elasticity of exports to the exporter-price index  $(\theta_i)$  – see eq. (3.15). As in section 3.2, proxy  $\theta_i$  using the exporter average bilateral trade elasticity.<sup>36</sup> Once  $\theta_i$  calibrated for each exporter, we can derive  $\zeta_j$  for each importer with  $\zeta_j = \frac{\theta_i\zeta_j}{\theta_i}$ ,<sup>37</sup> and create the exporter-importer-time-specific variable  $(1 - \zeta_j)\ln(w_{it})$ . To be consistent with the theory, we do not allow the PPML to reestimate  $\theta_i$  in eq. (3.19), but constrain  $\theta_i$  to take the values we estimated to derive  $\zeta_j$ .

$\beta$  is the average effect of RTAs. We find that  $\beta = 0.338$  (see appendix table 3.E.1). RTAs affecting trade through a reduction of trade costs ( $\tau$ ), the elasticity of exports to RTAs is subject to the bilateral trade elasticity distortion (eq. 3.15). Formally,  $\beta$  equals the elasticity of trade costs to RTAs time the average elasticity of exports to trade costs:  $\beta = \varepsilon_\tau(RTA) * (\overline{-\theta_i\zeta_j})$ . Therefore, to take into account the bilateral trade elasticity distortion and derive the exporter-importer-specific effect of RTAs, we simply need to replace the average trade elasticity  $(\overline{-\theta_i\zeta_j})$  in  $\beta = \varepsilon_\tau(RTA) * (\overline{-\theta_i\zeta_j})$  by the exporter-

<sup>36</sup>Knowing that  $\bar{\zeta} = 1$ , we have  $\theta_i = \frac{\sum_j \theta_i \zeta_j}{j}$ . See footnote 26.

<sup>37</sup>Because we use a proxy for  $\theta_i$ , we face some errors in  $\zeta_j$ . Computing, we do not find a purely importer-specific  $\zeta_j$ . To address this imprecision, we use the importer average of  $\frac{\theta_i \zeta_j}{\theta_i}$  as our proxy of  $\zeta_j$ .

importer-specific trade elasticity ( $-\theta_i\zeta_j$ ), which we estimated in section 3.2 (see figure 3.2.2).<sup>38</sup> Thus, we have:  $\beta_{ij} = \varepsilon_\tau(RTA) * (-\theta_i\zeta_j) = \beta * (-\theta_i\zeta_j)/(-\overline{\theta_i\zeta_j})$ .<sup>39</sup>

Second, in our General Equilibrium system,  $\tau$ , the trade cost parameter does not only captures trade costs related to RTAs but also other ones such as distance or language. We need to calibrate this level of trade costs. To do so, we re-estimate eq. (3.19) but constrain now the RTA elasticity to take the bilateral values  $\beta_{ij}$  that we estimated. Then, we use the exporter-importer fixed effects ( $\chi_{ij}$ ) to calibrate the level of trade costs unrelated to RTAs.<sup>40</sup>

To assess the impact of RTAs on welfare, we develop a counter-factual approach. We create a counterfactual global economy for 2013<sup>41</sup> in which no RTA exists. Compared to the baseline (the observed 2013 global economy), we can deduce the welfare effect of the 2013 RTAs network.

Therefore, the third step consist in defining our baseline outward (eq. 3.16) and inward (eq. 3.17) multilateral resistances. The PPML estimator has the nice property of obtaining estimates of the fixed effects from gravity estimations that are perfectly consistent with the structural gravity terms (Fally, 2015). Hence, using the PPML estimator, we can recover empirical expressions of the multilateral resistance terms:

$$\widetilde{\Pi}_i = \left( \frac{Y_i X_0}{\widetilde{\chi}_i} \right)^{-\frac{1}{\theta_i}} \quad (3.20)$$

$$\widetilde{\Phi}_j = \left( \frac{X_j}{X_0 \widetilde{\chi}_j} \right)^{-\frac{1}{\theta_j}} \quad (3.21)$$

where  $X_0$  is the reference level of expenditures.<sup>42</sup> To recover the baseline importer ( $\widetilde{\chi}_j$ ) and exporter ( $\widetilde{\chi}_i$ ) we re-estimate eq. (3.19) on the year 2013 restricting the estimation with the bilateral elasticities of RTAs ( $\beta_{ij}$ ) and wages ( $\theta_i(1 - \zeta_j)$ ), and the level of invariant trade costs ( $\widetilde{\chi}_{ij}$ ) that we estimated previously. Then, using equations 20 and 21 we can derive from  $\widetilde{\chi}_j$  and  $\widetilde{\chi}_i$  the baseline multilateral resistance terms.<sup>43</sup>

Fourth, we estimate the counter-factual multilateral resistances. We define our counter-factual and the RTA dummy to 0 for all country pairs – meaning that no RTA exist in this counter-factual 2013 global economy. We re-estimate eq. (3.19) on year 2013 calibrating the bilateral elasticities of RTAs ( $\beta_{ij}$ ) and wages ( $\theta_i(1 - \zeta_j)$ ), and the level of invariant trade costs ( $\widetilde{\chi}_{ij}$ ), but with RTA=0 for all

<sup>38</sup>We can also use the MMQR estimator to directly estimate the bilateral distortion of RTAs' elasticity. Yet, this will also capture the correlation between the value of trade and the effect of RTAs on trade costs. Empirically, this increases the bilateral distortion (see figure 3.E.1).

<sup>39</sup>In section 3.2 we estimated  $\overline{\theta_i\zeta_j} = \overline{\theta_j} = 3.383$ .

<sup>40</sup>Here we assume that all trade costs that are unrelated to RTAs are time-invariant. Alternatively, we can complete the estimation with exporter-importer-time variables which should affect trade costs and include them in the calibration of the trade cost level. Furthermore, some exporter-importer fixed effects are dropped due to convergence issues. These missing effects are replaced by regressing for 2013 the estimates of exporter-importer fixed effects on gravity variables and country fixed effects.

<sup>41</sup>The last year we observe trade, RTAs, tariffs and wages.

<sup>42</sup>We need to normalize one multilateral resistance term so that  $\widetilde{\phi}_0 = 1$ . Importer fixed effect  $\widetilde{\chi}_0$  is interpretable as  $X_0$  while all other fixed effects must be interpreted relative to  $X_0$ . In our empirical application we choose to normalize Germany's importer multilateral resistance term.

<sup>43</sup> $\widetilde{\theta}_i$  and  $\overline{\theta_j}$  are the estimates of  $\theta_i$  and  $\theta_j$  from previous steps.

country-pairs. Using the estimate of the counter-factual fixed effect  $\widetilde{\chi}_j^c$  and  $\widetilde{\chi}_i^c$ , we can derive from eq. (3.20) and (3.21) the counter-factual multilateral resistances  $\widetilde{\Phi}_j^c$  and  $\widetilde{\Pi}_i^c$ .

Fifth, we determine the counter-factual expenditures  $\widetilde{X}_i^c$  and output  $\widetilde{Y}_i^c$ . The trade balance implies that  $\frac{w_i^c}{w_i} = \frac{Y_i^c}{Y_i}$  (Anderson et al., 2018; Arkolakis et al., 2012). Assuming that the share of the outputs converted in expenditures is constant we also have:  $\frac{w_i^c}{w_i} = \frac{X_i^c}{X_i}$ . Therefore, any change in expenditures and outputs can be derived from the change in wage. Combining equation (3.18) with equation (3.20), we obtain the following relation between the change in wage and fixed-effects:

$$\frac{w_i^c}{w_i} = \left( \frac{\exp(\chi_i^c)X_0}{\exp(\chi_i)X_0^c} \right)^{-\frac{1}{\theta_i}} \quad (3.22)$$

From the estimated counter-factual fixed effects, we can therefore determine the change in wage between the counter-factual and the baseline, and derive the counter-factual expenditures and output:  $\widetilde{X}_i^c = \frac{w_i^c}{w_i} X_i$ ,  $\widetilde{Y}_i^c = \frac{w_i^c}{w_i} Y_i$ .

Sixth, we must take into account that these variations in wages, multilateral resistances, expenditures and outputs trigger new changes in trade, which then re-affect wages, multilateral resistances, outputs, expenditures and so forth. Based on the gravity equation (eq. 3.15), we can translate the changes in wages, multilateral resistances, outputs and expenditures, triggered initially by a change in trade cost, into export variations:

$$\widetilde{X}_{ij}^c = \frac{\widetilde{\tau}_{ij}^{-\theta_i \zeta_j} w_{ij}^c \theta_i (1-\zeta_j)}{\widetilde{\tau}_{ij}^{-\theta_i \zeta_j} \widetilde{w}_{ij} \theta_i (1-\zeta_j)} \frac{\widetilde{Y}_i^c \widetilde{X}_j^c \widetilde{\Pi}_i^{-\theta_i} \widetilde{\Phi}_j^{-\theta_j}}{Y_i X_j \widetilde{\Pi}_i^c \widetilde{\Phi}_j^c} \quad (3.23)$$

With these new trade-flows values we re-iterate steps 3, 4, 5 and 6 until prices converge.<sup>44</sup> This process over, we have finally defined our counter-factual economy where no RTA exists.

### 3.3.2 Results

The counter-factual general equilibrium built, we can calculate the change in exports per country  $\left( \frac{\sum_j X_{ij} - \sum_j \widetilde{X}_{ij}^c}{\sum_j X_{ij}^c} \right)$  and the change in welfare:  $\frac{W_i}{W_i^c} = \frac{Y_i / \widetilde{\Phi}_j}{Y_i^c / \widetilde{\Phi}_j^c}$  due to the RTAs network. We report results in table 3.3.1. We compare the GE effects with bilateral trade elasticity framework, to the GE effects estimated with constant trade elasticity. Results with constant trade elasticity are computed using the standard method from Anderson and Yotov (2016); Anderson et al. (2018); Fally (2015).

Bilateral trade elasticities induce three distortions which drive our results:

- i) Because trade elasticities  $(\theta_i \zeta_j)$  are exporter-importer-specific all RTAs will not have the same effect.<sup>45</sup> Empirically, since trade elasticities are larger for country-pairs with low trade-flow values,

<sup>44</sup>We consider convergence when the difference in change in wage is inferior to 0.01 and its standard deviation too.

<sup>45</sup>We assume that RTAs induce the same trade cost reduction. This assumption can be removed easily. We do it to



these country-pairs experience a larger expansion of trade from RTAs. Similarly, from the perspective of the exporter, signing an RTA with a small economy induces stronger (relative) change in exports than signing an RTA with a large economy.

ii) The elasticity of exports to the importer price-index ( $\theta_j$ ) is importer-specific and empirically decreases with the size of the importer. As detailed in section 3.2, a lower  $\theta_j$  implies a stronger sensitivity of the price index to changes in bilateral prices. Thus, in large economies, a reduction of trade costs implies a stronger reduction in the price index and, therefore, larger welfare gains. Similarly,  $\theta_i$  in eq. (3.22), derived from eq. (3.18) is exporter-specific and decreases with the economic size of the exporter. In these equations,  $-\theta_i$  defines the sensitivity of the output-price-index ratio  $\frac{Y_i}{\Pi_i^{-\theta_i}}$  to a change in wage (and inversely with  $\frac{1}{-\theta_i}$ ). Hence, a lower  $\theta_i$  induces that a change in  $\frac{Y_i}{\Pi_i^{-\theta_i}}$ , captured by exporter fixed effects  $\chi_i$  (see eq. 3.22), is associated with a larger change in wage, in output ( $\widetilde{Y}_i^c = \frac{w_i^c}{w_i} Y_i$ ) and, therefore, in welfare.

iii) In the gravity equation, the general negative effect of wage is captured by the output-price-index ratio  $\frac{Y_i}{\Pi_i^{-\theta_i}}$ . Yet, there subsists the bilateral distortion with the elasticity  $\theta_i(1 - \zeta_j)$  (see eq. 3.15). In this elasticity, the value of  $\zeta_j$  is key. If  $\zeta_j > 1$ , the part of the wage effect which is not captured by fixed effects has a negative impact on trade. On the contrary, if  $\zeta_j < 1$ , the effect is positive and the importer distortion mitigates the overall negative impact of a rise in wage. Taking into account this distortion in the GE, estimated changes in wages adjust. If  $\zeta_j < 1$  ( $\zeta_j > 1$ ), exporter fixed-effects are more (less) affected by a change in trade costs, leading to larger (lower) estimated variation in wage (see eq. 3.22). Empirically, we observe that, indeed,  $\zeta_j \in ]0; 2[$ .  $\zeta_j$  decreases with the importer's economic size. Therefore, this third distortion increases the estimated welfare gains of countries signing RTAs with large partners (i.e. destinations with low  $\zeta_j$ ), while it reduces the gains of countries whose RTAs partners are small economies (i.e. destinations with large  $\zeta_j$ ). The economic meaning of  $\zeta_j$  depends on the theoretical model (see section 3.1). But in the gravity equation (eq. 3.7, 3.11 and 3.15), it always captures the importer-specific distortion of the trade sensitivity to marginal costs (wage and trade costs). In the appendix section 3.D, we further discuss the theoretical and empirical implications of omitting the bilateral elasticity of wages.

In table 3.3.1, the differences in estimated exports and welfare gains between the bilateral trade elasticity approach (column 2-3 & 7-8) and the constant trade elasticity one (column 4-5 & 9-10) are driven by a mix of these three distortions. Once considering bilateral trade elasticities, we observe, in general, smaller changes in welfare. Economies which, on the contrary, experience stronger welfare changes are in the large majority very small economies, like Malta, Hong-Kong or Singapore. This supports a prevalence, in our sample, of the first distortion, affecting the direct effect of RTAs on simplify the identification of bilateral elasticities consequences.

trade: because bilateral trade elasticities are larger (smaller) for small (large) economies, RTAs signed by small (large) economies have stronger (weaker) direct impact on trade.<sup>46</sup>

These conclusions show that the existence of bilateral trade elasticities has important implications for welfare evaluations. Using an adapted frame is therefore crucial to estimate the welfare impact of trade policies and its heterogeneity.

### 3.4 Conclusion

This chapter develops an extensive analysis of the bilateral trade elasticities introduction in trade theory. By adapting the great categories of modern trade models, we provide a flexible theoretical framework consistent with the existence of bilateral trade elasticities. In their presence, the share of expenditure on domestic goods and the elasticity of exports to the importer price-index are sufficient statistics to infer welfare gains from trade. Computing them, we raise important distortions underlying the necessity of accounting for the country-pair sensitivity to trade cost. We provide a detailed methodology to perform policy evaluations with bilateral trade elasticities. Applying this approach to the Regional Trade Agreements, we show that ignoring the bilateral dimension of trade elasticities can lead to serious misestimations of trade liberalization's impact on exports and welfare. In addition, the pregnancy and heterogeneity of bilateral trade elasticities invite further research at the sector and firm level to better understand whether the mechanisms at work rely on the demand or supply side or a composition of both.

Implications for policy-makers are also important. Namely, our results raise that, even at the macro level, the same trade policy can have noticeably different effects on exports and real revenues depending on the fundamental economic characteristics of countries and their partners. Who wants to forecast trade policies' consequences need therefore to account for this variety in sensitivity to trade costs.

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<sup>46</sup>In appendix section 3.D, we show that for some very small economies the third distortion, related the bilateral elasticity of wages, is also important. The wage of these economies with very small internal market is highly sensitive to changes in exports. To what extent a change in wage affects in return trade is therefore crucial for them.

Table 3.3.1: Exports and welfare gains with bilateral trade elasticities

Trade elasticity:		Bilateral			Constant			Bilateral			Constant			
Country (Iso3)	Exports	Real revenues	Exports	Real revenues	Country (Iso3)	Exports	Real revenues	Exports	Real revenues	Country (Iso3)	Exports	Real revenues	Exports	Real revenues
ARG	8.68	1.69	16.18	0.60	ISR	12.92	3.43	14.65	3.71					
AUT	14.68	5.19	17.21	5.65	ITA	15.10	1.80	17.77	1.91					
BEL	2.48	12.02	2.67	13.09	JAM	1.54	3.81	1.47	4.46					
BGR	16.30	5.50	18.24	5.84	JPN	2.18	0.08	2.23	0.08					
BOL	10.35	0.23	11.21	0.02	KHM	11.65	1.87	22.27	0.44					
BRA	1.03	0.01	-0.23	-0.04	MLT	22.85	1.13	33.56	0.15					
CHN	2.42	0.06	2.22	0.05	MUS	0.34	-0.70	0.22	-0.82					
CRI	13.49	2.40	10.98	0.79	NLD	12.52	4.30	15.16	4.36					
CYP	47.60	0.90	57.98	0.08	NOR	17.03	2.34	19.24	2.49					
DEU	11.50	2.45	13.51	2.58	PAK	6.46	0.77	9.33	0.34					
ECU	-1.40	-0.36	-1.93	-0.43	PAN	15.21	0.38	17.33	0.04					
EGY	23.49	1.20	26.15	1.28	PHL	7.74	4.19	8.64	4.65					
ESP	17.22	2.51	20.40	2.69	POL	18.33	3.36	21.62	3.65					
FIN	14.33	2.14	16.17	2.27	PRT	19.49	4.14	22.74	4.56					
FRA	15.37	2.99	18.13	3.17	QAT	0.39	-0.48	0.24	-0.59					
GBR	14.76	2.52	17.41	2.69	ROM	16.87	5.58	19.66	6.09					
GHA	0.83	-1.46	1.09	-1.85	SAU	-1.64	-0.47	-2.01	-0.52					
GRC	24.86	2.24	27.95	2.39	SGP	10.34	3.57	13.71	2.94					
GTM	34.77	1.17	47.12	0.13	SLV	37.39	1.40	54.81	0.14					
HKG	23.20	2.85	34.88	1.78	THA	3.82	7.47	4.27	7.99					
HUN	9.65	7.99	11.17	8.79	URY	7.86	2.43	8.16	2.29					
IDN	14.29	2.08	16.91	2.21	USA	-0.88	-0.11	-1.56	-0.13					
IRL	5.29	6.92	5.95	7.61	VNM	6.90	4.63	8.02	5.03					
ISL	9.14	7.46	9.84	8.14	ZAF	10.37	1.03	10.67	1.41					

Notes: The real revenue is our measure of welfare. All numbers are variations in percentage.

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# Appendix

## 3.A Models complete derivation

Below, we present the full derivations of our models with bilateral trade elasticities.

### 3.A.1 Armington model derivation

We assume the following utility function with the bilateral elasticity of substitution  $\theta_i\zeta_j$ :

$$U = \int_{\Omega} q(\omega)^{\frac{\theta_i\zeta_j-1}{\theta_i\zeta_j}} d\omega \quad (3.24)$$

Agents maximise their utility under their revenue constraint. Deriving the Lagrangien, we obtain the following quantity/price ratio between varieties from country  $i$  and varieties from country  $l$ :

$$\frac{\frac{\theta_i\zeta_j-1}{\theta_i\zeta_j} q_{ij}^{\frac{\theta_i\zeta_j-1}{\theta_i\zeta_j}}}{\frac{\theta_l\zeta_j-1}{\theta_l\zeta_j} q_{lj}^{\frac{\theta_l\zeta_j-1}{\theta_l\zeta_j}}} = \frac{\lambda p_{ij}}{\lambda p_{lj}} \quad (3.25)$$

or

$$\frac{q_{ij}}{q_{lj}} = \frac{(\frac{\theta_i\zeta_j-1}{\theta_i\zeta_j} \lambda p_{ij})^{-\theta_i\zeta_j}}{(\frac{\theta_l\zeta_j-1}{\theta_l\zeta_j} \lambda p_{lj})^{-\theta_l\zeta_j}} \quad (3.26)$$

with  $q$  the quantity,  $p$  the price and  $\lambda$  the Langragien parameter. Considering a continuum of origins we therefore have:

$$\int_l p_{lj} q_{lj} = q_{ij} \left( \frac{\theta_i\zeta_j-1}{\theta_i\zeta_j} \lambda p_{ij} \right)^{-\theta_i\zeta_j} \int_l \left[ \left( \frac{\theta_l\zeta_j-1}{\theta_l\zeta_j} \lambda \right)^{-\theta_l\zeta_j} p_{lj}^{1-\theta_l\zeta_j} \right] \quad (3.27)$$

Defining  $X_j = \int_l p_{lj} q_{lj}$  the total demand in country  $j$ , and  $X_{ij} = q_{ij} p_{ij}$  the value of demand in  $j$  allocated to varieties produced in  $i$ , we have at the country-pair level:

$$X_{ij} = X_j \left( \frac{\theta_i\zeta_j}{\theta_i\zeta_j-1} \lambda \right)^{-\theta_i\zeta_j} p_{ij}^{1-\theta_i\zeta_j} \Phi_j^{\theta_j-1} \quad (3.28)$$



with the Dixit-Stiglitz price index:

$$\Phi_j = \left[ \int_l \left( \frac{\theta_l \zeta_j}{\theta_l \zeta_j - 1} \lambda \right)^{-\theta_l \zeta_j} p_{lj}^{1-\theta_l \zeta_j} \right]^{\frac{1}{1-\theta_j}} \quad (3.29)$$

Firms produce using only labour, with constant return to scale and homogeneous productivities normalised at 1. Having a monopole on their variety, they face the monopolistic competition price  $p_{ij} = \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} w_i \tau_{ij}$ ,  $w$  the wage and  $\tau$  the trade cost. Combining  $p_{ij}$  with equation (3.28), we obtain the gravity equation:

$$X_{ij} = X_j \frac{\left( \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \right)^{1-2\theta_i \zeta_j} \lambda^{-\theta_i \zeta_j} w_i^{1-\theta_i \zeta_j} \tau_{ij}^{1-\theta_i \zeta_j}}{\Phi_j^{1-\theta_j}} \quad (3.30)$$

Knowing from market clearance:  $Y_i = \sum_j X_{ij} = \sum_j X_j \left( \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \right)^{-\theta_i \zeta_j} p_{ij}^{1-\theta_i \zeta_j} \Phi_j^{\theta_j - 1} = \lambda^{-\theta_i} w_i^{1-\theta_i} \Pi_i^{1-\theta_i}$  that  $Y_i \Pi_i^{\theta_i - 1} = \lambda^{-\theta_i} w_i^{1-\theta_i}$ , we can make visible output  $Y_i$  and outward multilateral resistance  $\Pi_i$ :

$$X_{ij} = \frac{Y_i}{\Pi_i^{1-\theta_i}} \frac{X_j}{\Phi_j^{1-\theta_j}} \left( \frac{\theta_i \zeta_j}{\theta_i \zeta_j - 1} \right)^{1-2\theta_i \zeta_j} \lambda^{\theta_i(1-\zeta_j)} w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{1-\theta_j \zeta_i} \quad (3.31)$$

### 3.A.2 Heterogeneous firms model derivation

First steps of the heterogeneous firms model follow the same derivation as the previous model but keep assuming a standard Constant Elasticity of Substitution (CES) utility function  $U = \left[ \int_{\Omega} q(\omega)^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}} d\omega$  with  $\sigma$  the elasticity of substitution parameter. Deriving the consumer's utility maximisation problem we find that  $x_{ij}(\alpha)$  the value of country  $j$  demand allocated to the firm producing in country  $i$  with the marginal cost  $\alpha$  is:

$$x_{ij}(\alpha) = X_j \left( \frac{p_{ij}(\alpha)}{\Phi_j} \right)^{1-\sigma} \quad (3.32)$$

Firms' productivity  $\gamma = \frac{1}{\alpha}$  is heterogeneous (Chaney, 2008; Helpman et al., 2008). We assume firms in country  $i$  draw a marginal cost  $\alpha$  for each destination following the country-pair-specific inverse Pareto distribution:

$$G_{ij}(\alpha) = (\alpha^{\theta_i \zeta_j}) - \underline{\alpha}^{\theta_i \zeta_j} / \overline{\alpha}^{\theta_i \zeta_j} - \underline{\alpha}^{\theta_i \zeta_j} = \frac{\alpha^{\theta_i \zeta_j}}{\overline{\alpha}^{\theta_i \zeta_j}} \quad (3.33)$$

with  $\overline{\alpha} = \frac{X_j}{\Phi_j} \bar{\alpha}$  the maximum marginal cost,  $\underline{\alpha}$  the minimum one. For solving issues we follow Helpman et al. (2008) and assume that  $\underline{\alpha} = 0$ , i.e. that there is always at least one firm productive enough to export even a very small value. Aggregating sales at the country-pair level we therefore have:

$$X_{ij} = X_j \frac{N_i \int_0^{\alpha^*} p_{ij}(\alpha)^{1-\sigma} dG_{ij}}{\sum_l N_l \int_0^{\alpha^*} p_{lj}(\alpha)^{1-\sigma} dG_{lj}} \quad (3.34)$$

with  $N$  the number of firms and  $\alpha^*$  the marginal cost at which the firm is indifferent from selling in  $j$  from  $i$  or not. Removing from the integrals all parameters of the price function  $p_{ij} = \frac{\sigma}{\sigma-1} \alpha_{ij} w_i \tau_{ij}$  which do not depend on  $\alpha$  we can reshape the equation so that:

$$X_{ij} = X_j \frac{N_i (w_i \tau_{ij})^{1-\sigma} V_{ij}}{\sum_l N_l (w_l \tau_{lj})^{1-\sigma} V_{lj}} \quad (3.35)$$

with  $V_{ij} = \int_0^{\alpha^*} \alpha^{1-\sigma} dG_{ij}$ . Solving integrals we have:

$$V_{ij} = \frac{1}{1-\sigma} \bar{\alpha}^{-\theta_i \zeta_j} (\sigma^{\frac{\sigma}{1-\sigma}} (\sigma-1))^{\theta_i \zeta_j + 1 - \sigma} \left( \frac{X_j}{\Phi_j} \right)^{\frac{1}{\sigma-1} - 1} f_{ij}^{-\frac{\theta_i \zeta_j}{\sigma-1} - 1} (w_i \tau_{ij})^{\sigma-1-\theta_i \zeta_j} \quad (3.36)$$

where  $f_{ij}$  is the fixed trade cost from  $i$  to  $j$ . Finally, combining equations (3.34) and (3.35) we have the gravity equation:

$$X_{ij} = N_i \bar{\alpha}_i^{-(\theta_i \zeta_j)} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_i \zeta_j} w_i^{-(\theta_i \zeta_j)} \frac{X_j}{\Phi_j^{-\theta_j} \tau_{ij}^{-(\theta_i \zeta_j)} f_{ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1} - 1]}} \quad (3.37)$$

with

$$\Phi_j = \left( \sum_l N_l \bar{\alpha}_l^{-\theta_l \zeta_j} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_l \zeta_j} w_l^{-\theta_l \zeta_j} \tau_{lj}^{-\theta_l \zeta_j} f_{lj}^{-[\frac{\theta_l \zeta_j}{\sigma-1} - 1]} \right)^{-\frac{1}{\theta_j}} \quad (3.38)$$

or with  $\Pi_i$  the outward multilateral resistance and knowing from the market clearance condition that  $\frac{Y_i}{\Pi_i^{-\theta_i}} = N_i \bar{\alpha}_i^{-\theta_i} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{-\theta_i} w_i^{-\theta_i}$ :

$$X_{ij} = \frac{Y_i}{\Pi_i^{-\theta_i}} \frac{X_j}{\Phi_j^{-\theta_j}} \bar{\alpha}_i^{-(\theta_i \zeta_j)} \left( \frac{\sigma^{\frac{\sigma}{\sigma-1}}}{\sigma-1} \right)^{\theta_i (1-\zeta_j)} w_i^{\theta_i (1-\zeta_j)} \tau_{ij}^{-\theta_i \zeta_j} f_{ij}^{-[\frac{\theta_i \zeta_j}{\sigma-1} - 1]} \quad (3.39)$$

### 3.B Ricardian model

In Ricardian frames à la [Eaton and Kortum \(2002\)](#), the trade elasticity  $\theta$  is directly derived from the Fréchet distribution of productivity  $F_{ij}(\gamma) = Pr[\Gamma_i \leq \gamma] = e^{-T_i \gamma^{-\theta}}$ , where  $T_i$  is the origin's technological level,  $\gamma$  the productivity, and  $\theta$  captures the heterogeneity of goods in the economy. Because the model's derivation highly relies on the assumption that  $\theta$  the heterogeneity of goods in the economy. In order to derive bilateral trade elasticities, we operate two separate modifications: i) we assume that the shape of the productivity distribution is destination-specific ( $\theta_j$ ) – because of destination market restrictions and preferences, goods are not identical across importers.  $\theta_j$  captures the heterogeneity of goods specific to the destination market.<sup>47</sup> ii) we assume that the technology allows firms to adjust their productivity. Facing a change in marginal cost (higher  $w_i$  or  $\tau_{ij}$ ), firms'

<sup>47</sup>In keeping with [Desai and Trivedi \(2014\)](#), this assumption adopts a consumer-based definition of goods heterogeneity. The latter does not results only from the products' characteristics but also from the consumer perception and demand. Goods heterogeneity, therefore, differs depending on consumption markets.

incentives for new labour organisation or innovation adoption rise, as well as their financial constraints. Consequently, firms adapt their productivity (Akçigit and Melitz, 2022; Holmes and Schmitz Jr, 2010). We translate this behaviour into the production function  $q = \gamma'_{ij} l = \gamma(w_i \tau_{ij})^{1-\zeta_i} l$ , where  $\zeta_i$  determines to what extent firms adapt their productivity to a change in marginal costs.<sup>48</sup>

Thus, assuming  $\frac{\delta \zeta_i}{\delta \frac{X_i}{\Phi_i^{-\theta_i}}} < 0$  means that high real output is associated with a better productivity adaptation to marginal cost stress – i.e. lower  $\zeta_i$ . Respectively,  $\frac{\delta \theta_j}{\delta \frac{Y_j}{\Pi_j^{-\theta_j}}} < 0$  means that high real expenditures are associated with a higher heterogeneity in goods – i.e. lower  $\theta_j$  (Bils and Klenow, 2001; Chai et al., 2023; Saviotti, 2002; of Labor Statistics, 2006).

From the productivity distribution  $F_{ij}(\gamma) = Pr[\Gamma_i \leq \gamma] = e^{-T_i \gamma^{-\theta_j}}$  and the production function  $q = \gamma(w_i \tau_{ij})^{1-\zeta_i} l$  we can therefore derive the price and aggregated price functions:

$$p_{ij} = (w_i \tau_{ij})^{\zeta_i} \alpha \quad \Phi'_j = \sum_{i=1}^N T_i (w_i \tau_{ij})^{-\theta_j \zeta_i} \quad (3.40)$$

where  $\alpha = \frac{1}{\gamma}$ . Solving the model, we obtain the gravity equation:

$$X_{ij} = X_j \frac{T_i (w_i \tau_{ij})^{-\theta_j \zeta_i}}{\Phi_j^{-\theta_j}} \quad (3.41)$$

where  $\Phi_j = \left( \sum_{i=1}^N T_i (w_i \tau_{ij})^{-\theta_j \zeta_i} \right)^{-\frac{1}{\theta_j}}$ . The bilateral trade elasticity is therefore  $-\theta_j \zeta_i$ . Assuming symmetry of elasticities  $\theta_j \zeta_i = \theta_i \zeta_j$ , we can rewrite the gravity equation and make visible the output  $Y_i$  and the outward multilateral resistance term  $\Pi$ :

$$X_{ij} = \frac{X_j}{\Phi_j^{-\theta_j}} \frac{Y_i}{\Pi_i^{-\theta_i}} w_i^{\theta_i(1-\zeta_j)} \tau_{ij}^{-\theta_j \zeta_i} \quad (3.42)$$

We observe the same three main changes as in the Armington and heterogeneous firms frameworks. i) The trade elasticity  $-\theta_i \zeta_j$  is exporter-importer-specific. ii) The elasticity of exports to the importer price-index ( $\Phi_j$ ) is importer-specific and the elasticity of exports to the exporter price-index ( $\Pi$ ) is exporter-specific. Again these elasticities are determined by  $\theta$ . In the Ricardian framework, this means that the exports' sensitivity to aggregated prices depends on the country-specific heterogeneity in goods. iii) In eq. (3.26), the overall negative effect of the wage toward all destinations is captured by the output-price-index ratio  $\left( \frac{Y_i}{\Pi_i^{-\theta_i}} \right)$ . Yet, there subsists the wage bilateral distortion  $(w_i^{\theta_i(1-\zeta_j)})$ , which increases or decreases the negative effect of the wage on bilateral exports depending on whether

<sup>48</sup>When  $\zeta_i < 1$  firms mitigate the effects of changes in marginal costs and improve their productivity. Yet, when  $\zeta_i > 1$  firms adapt but amplifying the effects of marginal costs changes. This reflects a supply structure where the financial difficulties induced by trade costs overcome the incentives for innovation, leading to a reorganisation impacting productivity negatively. Such amplifying behaviour is particularly characteristic of low-productivity economies (Akçigit and Melitz, 2022; Melitz and Trefer, 2012).

$\zeta_j$  is superior or inferior to one. In the Ricardian frame, this bilateral distortion of the wage effect on exports means that the wage has a more negative effect on exports toward destinations where the heterogeneity in goods (defined by  $\theta_j$ ) is the lowest.<sup>49</sup>

### 3.C Gains from trade, theoretical proof

In this section, we develop the proof of equation (3.15), defining welfare gains from trade. By definition  $dlnW_j = dln\frac{Y_j}{\Phi_j}$ . Noting that trade balance implies  $dlnY_j = dlnw_j$  and using labour in country  $j$  as our numeraire, we have  $dln(Y_j) = 0$ . Furthermore, despite price index  $\Phi_j$  is not equivalent in our three models, parameters making these differences are all constants (Arkolakis et al., 2012). Therefore, in all our models with bilateral trade elasticities, welfare gains from trade are defined by:

$$dlnW_j = -dln\Phi_j = -dln \left[ \left( \sum_{i=1}^n (w_i \tau_{ij})^{-\theta_i \zeta_j} \right)^{\frac{1}{-\theta_j}} \right] = - \sum_{i=1}^n \psi_{ij} \frac{-\theta_i \zeta_j}{-\theta_i} [dln(w_i) + dln(\tau_{ij})] \quad (3.43)$$

where  $\psi_{ij}$  is the market share of country  $i$  in country  $j$ . By the gravity equation, changes in relative imports are such that:

$$dln(\psi_{ij}) - dln(\psi_{jj}) = (-\theta_i \zeta_j)(dln(w_i) + dln(\tau_{ij})) \quad (3.44)$$

Combining equation (3.42) with equation (3.43) we obtain:

$$dlnW_j = - \sum_{i=1}^n \psi_{ij} \frac{1}{-\theta_i} [dln(\psi_{ij}) - dln(\psi_{jj})] \quad (3.45)$$

Knowing that  $\sum_{i=1}^n \psi_{ij} = 1$  we can derive:

$$dln(W_j) = \frac{dln(\psi_{jj})}{-\theta_j} \quad (3.46)$$

Finally, defining  $\hat{\mu} \equiv \mu'/\mu$  the change in any variable  $\mu$  between the initial and the new equilibrium, we find back equation (3.15):

$$\hat{W}_j = \hat{\psi}_{jj}^{\frac{1}{-\theta_j}} \quad (3.47)$$

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<sup>49</sup>Knowing we assumed that  $\theta_j \zeta_i = \theta_i \zeta_j$ , we have  $(1 - \theta_j) \zeta_i = \theta_i (1 - \zeta_j)$ .

### 3.D Bilateral wage elasticities omission

In section 3.1, we adapted the classical gravity models presented in (Arkolakis et al., 2012) to bilateral trade elasticities. In this category of models, such an adaptation also leads to the presence of bilateral wage elasticities. Our analysis mainly focuses on the impact of trade costs, yet wages and their elasticities play a central role in general equilibrium. In this section, we therefore discuss the theoretical and empirical implications of omitting the bilateral dimension of the wage elasticities.

#### 3.D.1 Theoretical implication

Having in structural gravity bilateral trade elasticities with exporter-specific wage elasticities requires that we introduce a second level of assumption in our models.

##### Armington model

In the Armington frame, the first level of assumption follows the same logic as in our main model. To observe an elasticity of trade dependent on the exporter we introduce in the utility function an exporter-specific elasticity of substitution  $\theta_i$ :

$$U = \int_{\Omega} q(\omega)^{\frac{\theta_i-1}{\theta_i}} d\omega \quad (3.48)$$

As in section 3.1, the consumer weights the goods differently depending on their origin. Yet, here this exporter-specific weights are the same for each destination. This assumption leads to exporter-specific trade and wage elasticities.

To observe bilateral trade elasticities in the gravity equation, we must add a second level of assumption on the supply side. We assume that firm can adapt their productivity to trade costs depending on the parameter  $\zeta_j'$ . We already proposed adaptive productivity assumption in section 3.B. Nevertheless, the main difference here is that we do not assume global productivity adaptation to marginal costs (Akcigit and Melitz, 2022; Holmes and Schmitz Jr, 2010), but only to trade costs. Therefore, the economic interpretation is no longer about innovation and productivity improvement to face a marginal cost stress, but a productivity selection depending on the destination market's importance that  $1 - \zeta_j$  captures.<sup>50</sup> Thus, we can define the production function as follows:

$$q_{ij} = \tau_{ij}^{1-\zeta_j} l \quad (3.49)$$

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<sup>50</sup>If  $\zeta_j$  is low, destination  $j$  is an important market. The firm will highly react to marginal costs specific to this destination and allocate the most productive factors to the production of the goods intended for this destination.

Running through the model, we obtain the gravity equation:

$$X_{ij} = X_j \frac{\left(\frac{\theta_i}{\theta_i-1}\right)^{1-2\theta_i} \lambda^{-\theta_i} w_i^{1-\theta_i} \tau_{ij}^{(1-\theta_i)\zeta_j}}{\Phi_j^{1-\theta_j}} \quad (3.50)$$

where the trade elasticity  $(1-\theta_i)\zeta_j$  is bilateral and the wage elasticity  $1-\theta_i$  exporter-specific. Making visible the output  $Y_i$  and the outward multilateral resistances  $\Pi_i$  we have:

$$X_{ij} = \frac{Y_i}{\Pi_i^{1-\theta_i}} \frac{X_j}{\Phi_j^{1-\theta_j}} \tau_{ij}^{(1-\theta_i)\zeta_j} \quad (3.51)$$

The wage and its elasticity being now both only exporter-specific, they are captured by the output/price ratio  $\frac{Y_i}{\Pi_i^{1-\theta_i}}$ . Yet, this is at the cost of stronger assumptions on the micro-economic structure.

### Ricardian model

To derive bilateral trade elasticities, the Ricardian frame already required the existence of productivity adaptation to marginal costs (see section 3.B. So that exporter-specific wage elasticities and bilateral trade ones coexist, we need again to reach another assumption level. We define the following production function:

$$q = \gamma w_i^{1-\frac{\zeta_i}{\zeta_j}} \tau_{ij}^{1-\zeta_i} l \quad (3.52)$$

There, we still assume that marginal cost stress leads firms to innovation and productivity improvement (Akcigit and Melitz, 2022; Holmes and Schmitz Jr, 2010). Yet, wage being an exporter-level cost, here firms experience difficulties in adapting all production-chains depending on the destination economy's complexity defined by  $\zeta_j$ . Therefore, with this additional assumption, the parameter  $\zeta_j$  determines the capacity for economies to adapt to marginal costs as well as the adaptation restriction they impose on concurrent economies.

Running through the model, we derive the following gravity equation:

$$X_{ij} = X_j \frac{T_i w_i^{-\frac{\theta_j \zeta_i}{\zeta_j}} \tau_{ij}^{-\theta_j \zeta_i}}{\Phi_j^{-\theta_j}} \quad (3.53)$$

with

$$\Phi_j = \left[ \sum_{i=1}^N T_i w_i^{-\frac{\theta_j \zeta_i}{\zeta_j}} \tau_{ij}^{-\theta_j \zeta_i} \right]^{-\frac{1}{\theta_j}} \quad (3.54)$$

Assuming symmetry of elasticities  $\theta_i \zeta_j = \theta_j \zeta_i$  as in section 3.B we finally have:<sup>51</sup>

$$X_{ij} = X_j \frac{T_i w_i^{-\theta_i} \tau_{ij}^{-\theta_i \zeta_j}}{\Phi_j^{-\theta_j}} \quad (3.55)$$

where the wage elasticity  $\theta_i$  is exporter-specific and the trade elasticity  $-\theta_i \zeta_j$  is bilateral. Nonetheless, in the Ricardian frame this requires the highest level of assumption.

### Heterogeneous firms model

In our baseline model with heterogeneous firms, bilateral elasticities are derived from the Pareto shape parameters of the productivity distribution (see section 3.1.3). As in the Armington frame, to observe exporter-specific wage elasticities, we need to split the sources of heterogeneity in trade elasticities ( $i$  and  $j$ ). We therefore define the exporter-specific productivity distribution:

$$G_i(\alpha) = (\alpha^{\theta_i} - \underline{\alpha}^{\theta_i}) / (\bar{\alpha}_i^{\theta_i} - \underline{\alpha}^{\theta_i}) = \frac{\alpha^{\theta_i}}{\bar{\alpha}_i^{\theta_i}} \quad (3.56)$$

Again, we complete the first assumption level with the introduction of adaptive productivity. Firms improving their productivity in reaction to trade-cost stress, they face the production function:

$$q = \gamma \tau_{ij}^{1-\zeta_j} f_{ij}^{\zeta_j^f} l \quad (3.57)$$

where  $\zeta_j$  determines the degree of adaptation to variable trade costs and  $\zeta_j^f$  to fixed trade costs.<sup>52</sup> The economic interpretation of this assumption on productivity is the same as in the Armington case (see section 3.D.1).

From this new set-up we can derive the gravity equation:

$$X_{ij} = N_i \bar{\alpha}_i^{-\theta_i} w_i^{-\theta_i} \frac{X_j}{\Phi_j^{-\theta_j} \tau_{ij}^{-(\zeta_j \theta_i)}} f_{ij}^{-[\frac{\zeta_j^f \theta_i}{\sigma-1} - \zeta_j^f]} \quad (3.58)$$

Variable and fixed trade elasticities  $\zeta_j \theta_i$  and  $\frac{\zeta_j^f \theta_i}{\sigma-1} - \zeta_j^f$  are bilateral, while the wage elasticity is exporter-specific. Yet, again this requires specific assumptions on productivity and firms' behaviour.

### 3.D.2 Policy evaluation

We replicate the approach presented in section 3.3.1 but with a general equilibrium structure derived from our models associating bilateral trade elasticities and exporter-specific wage elasticities.

<sup>51</sup>Under this assumption we have  $-\frac{\theta_j \zeta_i}{\zeta_j} = -\frac{\theta_i \zeta_j}{\zeta_i} = -\theta_i$

<sup>52</sup>We make no assumption on the relation between  $\zeta_j$  and  $\zeta_j^f$ .

The GE is characterized by the four following equations:

$$X_{ij} = \frac{Y_i}{\Pi_i^{-\theta_i}} \frac{X_j}{\Phi_j^{-\theta_j}} \tau_{ij}^{-\theta_i \zeta_j} \quad (3.59)$$

$$\Pi_i^{-\theta_i} = \sum_j \frac{\tau_{ij}^{-\theta_i \zeta_j}}{\Phi_j^{-\theta_j}} X_j \quad (3.60)$$

$$\Phi_j^{-\theta_j} = \sum_i \frac{\tau_{ij}^{-\theta_i \zeta_j}}{\Pi_i^{-\theta_i}} Y_i \quad (3.61)$$

$$w_i = \left( \frac{Y_i}{\Pi_i^{-\theta_i}} \right)^{-\frac{1}{\theta_i}} \quad (3.62)$$

The only difference with the structure presented in section 3.3.1 is the absence of the bilateral wage elasticities. Empirically, we follow the same process as previously without specifying wages, since with this GE structure they are fully captured by fixed effects. We report the policy evaluation results in table 3.D.1 (columns 4-5 & 9-10). We compare exports and welfare changes to our baseline results (columns 2-3 & 7-8). For the majority of countries, the difference is marginal. Yet very small economies, like Hong-Kong, Singapore or Malta, are particularly sensitive to the inclusion or omission of bilateral wage elasticity. Signing RTAs with large economies (i.e. with low  $\zeta_j$ ), the negative effect of wage on their exports is mitigated, inducing larger change in real revenue (see section 3.3).

Thus, for the majority of countries, including bilateral wage elasticities poorly affects policy evaluation results. Yet, for some specific countries, which may change depending on the evaluated policy, this can lead to very different welfare estimates. Moreover, as demonstrated in our theoretical discussion, associating bilateral trade elasticities with exporter-specific wage elasticities requires stronger assumptions on productivity and agents' behaviour.



Table 3.D.1: Exports and welfare gains with bilateral trade elasticities &amp; constant wage elasticity

Trade elasticity:		Bilateral			Constant			Country (Iso3)			Bilateral			Constant		
Country (Iso3)	Exports	Real revenues	Exports	Real revenues	Exports	Real revenues	Exports	Real revenues	Country (Iso3)	Exports	Real revenues	Exports	Real revenues	Exports	Real revenues	
ARG	13.98	0.55	16.18	0.60	ISR	12.82	3.42	14.65	3.71							
AUT	14.69	5.20	17.21	5.65	ITA	15.12	1.81	17.77	1.91							
BEL	2.45	12.07	2.67	13.09	JAM	1.43	4.23	1.47	4.46							
BGR	16.40	5.47	18.24	5.84	JPN	2.03	0.08	2.23	0.08							
BOL	10.51	0.02	11.21	0.02	KHM	19.99	0.40	22.27	0.44							
BRA	-0.04	-0.04	-0.23	-0.04	MLT	31.60	0.14	33.56	0.15							
CHN	1.98	0.05	2.22	0.05	MUS	0.42	-0.65	0.22	-0.82							
CRI	9.61	0.69	10.98	0.79	NLD	12.99	4.08	15.16	4.36							
CYP	53.78	0.08	57.98	0.08	NOR	16.99	2.33	19.24	2.49							
DEU	11.53	2.45	13.51	2.58	PAK	8.26	0.31	9.33	0.34							
ECU	-1.53	-0.36	-1.93	-0.43	PAN	15.98	0.04	17.33	0.04							
EGY	23.76	1.22	26.15	1.28	PHL	7.50	4.17	8.64	4.65							
ESP	17.29	2.52	20.40	2.69	POL	18.38	3.36	21.62	3.65							
FIN	14.31	2.13	16.17	2.27	PRT	19.42	4.17	22.74	4.56							
FRA	15.41	3.00	18.13	3.17	QAT	0.39	-0.47	0.24	-0.59							
GBR	14.83	2.53	17.41	2.69	ROM	16.93	5.55	19.66	6.09							
GHA	1.08	-1.56	1.09	-1.85	SAU	-1.55	-0.44	-2.01	-0.52							
GRC	24.83	2.24	27.95	2.39	SGP	11.70	2.73	13.71	2.94							
GTM	41.21	0.11	47.12	0.13	SLV	47.74	0.13	54.81	0.14							
HKG	29.22	1.63	34.88	1.78	THA	3.70	7.43	4.27	7.99							
HUN	9.70	7.97	11.17	8.79	URY	7.50	2.14	8.16	2.29							
IDN	14.39	2.04	16.91	2.21	USA	-1.18	-0.12	-1.56	-0.13							
IRL	5.29	6.96	5.95	7.61	VNM	6.92	4.61	8.02	5.03							
ISL	9.16	7.43	9.84	8.14	ZAF	9.63	1.31	10.67	1.41							

Notes: The real revenue is our measure of welfare. All numbers are variations in percentage.

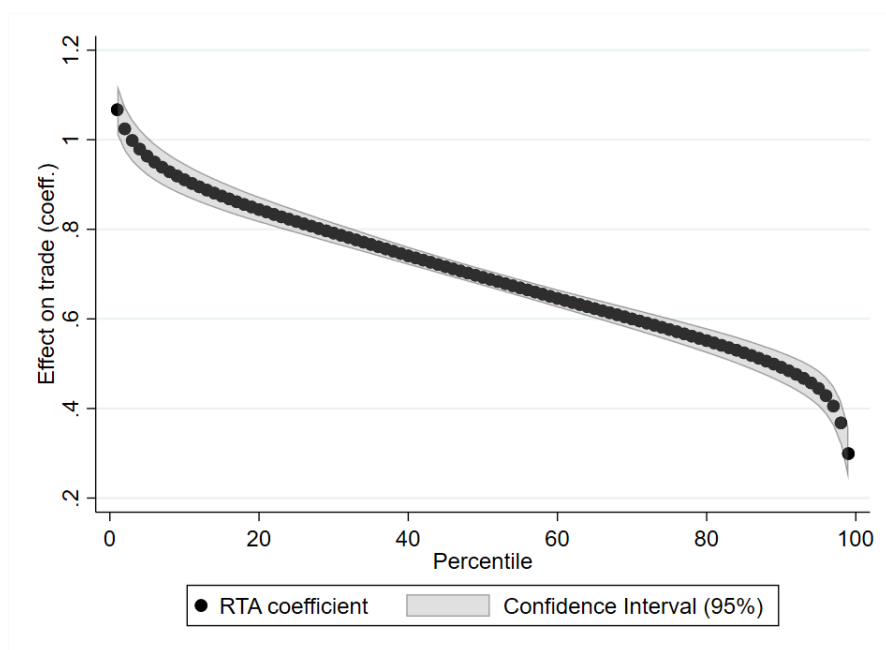
### 3.E Supplementary tables and figures

Table 3.E.1: Estimating the average trade elasticity

Estimator: PPML		
Dependent variable	Bilateral exports	Bilateral exports
Variables	(1)	(2)
$\ln(1+\text{Tariffs})$	-3.383 <sup>a</sup> (0.478)	
RTA		0.338 <sup>a</sup> (0.021)
Exporter x Year FE	yes	yes
Importer x Year FE	yes	yes
Exporter x Importer FE	yes	yes
No. observ.	51.140	303.300

*Note:* PPML, Poisson Pseudo Maximum Likelihood; FE, Fixed effects; Dependent variable is exports from country  $i$  to country  $j$  at time  $t$  in millions of current dollars. Standard errors clustered at country-pair level are in parentheses. a, b, c denote significantly different from 0 at 1%, 5% and 10% level, respectively.

Figure 3.E.1: Quantile estimates of RTAs



*Note:* Estimator is a Method of Moments-Quantile Regression with exporter-year and importer-year fixed effects.

# General conclusion

This thesis aims to contribute to the understanding of the interplay between international insecurity and economic activities. It focuses on the economic costs of insecurity, states' political and cooperative responses, and differences in sensitivity to exchange costs.

The first finding of this thesis is the identification of efficient tools for states to cooperate in reducing international insecurity and its adverse economic effects. The first chapter emphasizes the significant role of military alliances and their profound impact on global trade. It underscores the necessity of forging deep alliances with economic and political partners, highlighting the substantial economic returns that both developed and developing countries would accrue from expanding such agreements, particularly benefiting small economies. Furthermore, at the microeconomic level, the second chapter concludes that direct foreign interventions also enhance security and support local firms' activities. However, it raises the importance of cooperation with the local state and the imperative of focusing on long-term peace and state-building efforts to observe positive outcomes.

Conflicts are multifaceted. This thesis demonstrates their diversity and varying economic effects depending on their characteristics. Actors, timing, intensity, location, exposure – the context is key to understand their dynamic. Additionally, security policies hinge not on the absence of violence but on its management. To ensure the protection of agents, prevent significant material damages, suppress armed threats, or more generally enforce the rule of law and the security conditions economic activities require, these policies may result in violent events at times. In the first chapter, I demonstrated that alliances reduce international insecurity but increase the occurrence of cooperative military events between partners. Similarly, in the second chapter, I show that in the state-building context of the 2003-2016 Afghan conflict, exposure to military events had, on average, a positive effect on formal firms. While some events signify pure insecurity, others, beyond direct damages, may convey an expansion of state capacity and contribute to peace-building and international security. Therefore, analysing conflicts necessitates considering this heterogeneity to understand their economic impact and devise appropriate policies.

Sensitivity to insecurity is largely heterogeneous. At the micro-level, the second chapter shows

that insecurity shocks affect small firms more. At the macro-level, the first chapter demonstrates that small countries gain proportionally more from signing military alliances. The third chapter directly investigates this issue and proposes a flexible framework with country-pair-specific sensitivity to trade costs applicable to international insecurity. Upon considering these heterogeneities, it shows that all countries would benefit from the development of a common cooperative security system. Nonetheless, a significant disparity exists between interest and political capacity to promote such cooperation. Small economies would be the primary beneficiaries but are less integrated into the security cooperation network. Conversely, large economies are deeply integrated and capable of advocating for the global expansion of security cooperation through direct interventions or treaties like alliances. However, they are less vulnerable to international insecurity and thus have less economic interest in reducing geopolitical tensions and fostering collective security cooperation.

Economic activities and trade are inherently linked to the management of violence. Many research questions remain open. To better understand this relationship, future work should delve into the complexity of international insecurity and conduct precise evaluations of decades of security policies. Moreover, considering the diversity of interests and capacities, and the strength of geopolitical tensions, these concluding remarks call for involvement from international organizations. In the past, they succeeded in uniting countries through collective trade liberalization. To take a new step toward peace and globalization, their involvement in promoting security cooperation will be crucial.



## RÉSUMÉ

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Les activités économiques et le commerce sont profondément liés à la gestion de la violence, constituant un défi majeur dans le système mondial moderne. Cette violence, englobant l'expropriation, la destruction, la piraterie, le terrorisme, les conflits et les guerres, affecte tous les pays. Y être exposé entraîne des coûts économiques substantiels, entravant le commerce et le développement. En réponse, les États ont mis en place des politiques de sécurité et attendent en retour des bénéfices économiques. Cette thèse analyse la connexion entre l'insécurité internationale et les activités économiques. Le premier chapitre se concentre sur l'effet des alliances militaires sur le commerce - des traités spécifiquement conçus pour réduire l'insécurité internationale. Dans le deuxième chapitre, nous examinons la complexité des événements militaires et étudions leur impact micro-économique sur les entreprises formelles. Enfin, le chapitre trois analyse les conséquences de sensibilités bilatérales aux coûts d'échange, y compris l'insécurité, sur les revenus réels.

## MOTS CLÉS

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Insécurité internationale, Gravité Structurale, Commerce, Alliances Militaires, Afghanistan, Firmes Formelles, Elasticités Hétérogènes, Equilibre Général, Bien-Etre

## ABSTRACT

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Economic activities and trade are deeply intertwined with the management of violence, posing a significant challenge in the modern global system. This violence, spanning expropriation, destruction, piracy, terrorism, conflicts, and wars, affects all countries. The exposure to violence leads to substantial economic costs, hindering trade and development. In response, states have enforced security policies and expect economic benefits in return. The thesis analyses this connexion between international insecurity and economic activities. The first chapter focuses on the military alliances' effect on trade – treaties specifically designed to reduce international insecurity. In the second chapter, we dig into the complexity of military events and investigate their micro-economic impact on formal firms. Finally, chapter three analyses the consequences of country-pair-specific sensitivity to exchange costs, including insecurity, on real revenues.

## KEYWORDS

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International Insecurity, Structural gravity, Trade, Military Alliances, Afghanistan, Formal Firms, Heterogeneous Elasticities, General Equilibrium, Welfare